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ABSTRACT

This is the first of two volumes of selected papers presented at the 1996 conference on "Changes in Working Time in Canada and the United States." Eleven chapters focus on weekly hours worked by individuals, including the recent changes in the distribution of weekly working time in Canada and the U.S., implications of the changing distribution of hours worked for earnings inequality, and efforts to reduce unemployment through mandated hours reduction. "Introduction and Overview" (Ging Wong, Garnett Picot) is followed by these three chapters in Part I that offer statistical overviews and analyses of the trends in working hours for Canada and the U.S.: "The Changing Workweek: Trends in Weekly Hours of Work in Canada, 1976-95" (Mike Sheridan et al.); "Trends in Hours of Work in the U.S." (Philip L. Rones et al.); and "Working Hard" (Linda Bell, Richard B. Freeman). Three chapters in Part II consider the role of working time on increased earnings inequality: "Working Time, Wages, and Earnings Inequality Among Men and Women in Canada, 1981-93" (Garnett Picot); "Patterns of Foregone Potential Earnings Among Working-Age Males, 1975-92" (Robert Haveman et al.); and "The Growth of Income and Employment Inequality in Australian Cities" (R.G. Gregory, Boyd Hunter). These two chapters in Part III examine issues related to labor supply and hours constraints: "Supply of Hours Per Day and Days Per Week--Evidence from the Canadian Labor Market Activity

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Survey" (Richard E. Mueller) and "Hours Constraints: Theory, Evidence, and Policy Implications" (Kevin Lang, Shulamit Kahn). These three chapters on Part IV focus on "short-time compensation (STC) programs": "Measuring the Effects of STC on Workforce Dynamics" (Karen Needels, Walter Nicholson); "Short-Time Work in the U.S.: Implications for Evaluation of STC Schemes" (Alec R. Levenson); and "Economic Activities and the Demand for Work Sharing in Canada" (Tim Siedule et al.). Cited author and subject indexes are appended. Each paper contains references. (YLB)

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Working Time in Comparative Perspective. Volume I: Patterns, Trends, and the Policy Implications for Earnings Inequality and Unemployment

Ging Wong and Garnett Picot, Editors

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Editors

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and
Garnett Picot
Editors

2001

W.E. Upjohn Institute for Employment Research
Kalamazoo, Michigan

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8

Introduction and Overview

Ging Wong
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This is the first of two volumes of selected papers presented at the conference on "Changes in Working Time in Canada and the United States," which was held in Ottawa, Ontario, on June 13–15, 1996, and was jointly sponsored by the Canadian Employment Research Forum (CERF), the W.E. Upjohn Institute for Employment Research, and Statistics Canada. It reflects a renewed interest in recent years in the empirical evidence for changing labor supply—both hours of work and labor market participation—and the implications for employment, income support benefits, and taxation policies and programs.

To place this policy and research issue in a Canadian context, a February 1995 Parliamentary Committee report on Social Security reform called for initiatives to better understand and make policy recommendations regarding the redistribution of working time. What was clear to policymakers was that employment and income security policies, including unemployment insurance and welfare reform, needed to address significant changing patterns of work arrangements. At the same time, policymakers were handicapped by the absence of a knowledge base in this area. The research literature made sporadic references to the growth of a contingent workforce, flexible working arrangements, and nonstandard employment, but there was a lack of focus that pulled together the relevant concepts and empirical evidence.

This myopia is understandable. Through the 1980s, hours of work received little attention by academics and policy analysts, at least compared with issues such as unemployment (in Canada) and wage inequality (in North America). Average hours worked had declined slowly over many decades and then stabilized. This stasis produced little excitement. Labor supply topics received considerable attention during the 1970s, but this work also waned through the 1980s.

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During the 1990s, however, researchers observed that although average hours worked changed little, the manner in which the economy distributed working time was being transformed. Hours were becoming more polarized; some workers were working more hours, others fewer. This held implications for earnings inequality, unemployment and underemployment (for some), and overwork (for others). Many began to wonder what had happened to the promise of increased leisure. Paradoxically, within the context of high unemployment in Canada and Europe, many analysts were concerned about overwork and increased time-stress among a significant portion of the population. Work sharing or short-time compensation was seen as one response to the unemployment problem in Europe and Canada.

Issues regarding hours of work among women in particular also materialized. Often related to the conflict between work and family responsibilities, these concerns led to research on "flexible hours," job sharing, and other ways of providing increased flexibility in hours. The impact on worker performance and firm productivity of this potentially increased working-time flexibility also came to the fore. These and other events refocused the limelight on working time and resulted in the conference from which chapters in these volumes were selected.

The conference was international in coverage, in recognition that the same economic trends leading to pressures for changing employment relationships were present in Canada, the United States, and other industrialized countries. Noted researchers such as Richard Freeman have identified this issue to be of central importance in understanding how labor markets are evolving. The International Labor Organization (ILO) also recognized the transformations taking place in Michael White's *Working Hours: Assessing the Potential for Reduction.*[1] A more recent and more provocative piece in the U.S. literature is Jeremy Rifkin's *The End of Work.*[2]

The purpose of this book and the companion volume is to describe and place this transformation in a comparative and historical perspective, as well as to examine some of the new research and policy issues that have emerged in its wake. Most of the chapters in these volumes examine the situation in Canada or the United States, though some also look at working-time issues in western Europe and Australia. The essays in this volume present no central thesis, although there are recurring themes. It should be noted that the original conference papers have been revised for this collection.

The three chapters in Part I offer statistical overviews and analyses of the trends in working hours for Canada and the United States. Michael Sheridan, Deborah Sunter, and Brent Diverty provide empirical evidence on the Canadian trends in weekly hours of work from 1976 to 1995. Drawing on data from the monthly Labour Force Survey (LFS) and the Survey of Labour and Income Dynamics, they quantify the perceived shift away from the standard workweek (35–40 hours). Specifically, the analysis assesses movements in weekly hours as they relate to key labor market indications, including class of work, age, sex, education, occupation, and industry. The data support the concept that, while working hours are in flux, the distribution of hours has polarized over time. Evidence from the last 20 years suggests that hours have shifted from standard to both long (41 hours or more) and short workweeks (34 hours or less). The shift to shorter workweeks appears to have been triggered by the recessions of both the early 1980s and 1990s, while most of the growth in the share of long working time appears to be a more recent occurrence. Over the period, the 10.7-percentage-point decline in the proportion of people working standard hours was coincident with increases of 7.6 percentage points in short hour weeks and 3 percentage points in long hours. Further, this primarily unidirectional shift toward the hours distribution poles is evident when workers are grouped by various characteristics. The hours polarization persists, though to a lesser extent, when special groups such as the self-employed, multiple jobholders, young workers, and managers are removed from the analysis.

The chapter by Philip Rones, Jennifer Gardner, and Randy Ilg supplies a parallel analysis of trends in hours of work in the United States. Like Canada, the United States has experienced a sizable increase in the fraction of workers working very long hours, though, unlike Canada, the United States has not seen an increase in the fraction working short hours. The growth in the share of workers working 49 or more hours per week accounted for a modest increase in average weekly hours worked from 1976 to 1993. Interestingly, growth in the fraction working very long hours occurred among both men and women and within all major occupational categories.

Even more striking has been the growth in annual hours worked by women since the 1970s. From 1976 to 1993, average annual working hours among prime-age working women (25–54 years old) increased by 45 percent. This large increase was due both to the increase in labor force participation among women and to the fact that, when they join the labor force, women are more likely to work year round. Average annual hours worked by men has changed little since the 1970s.

Linda Bell and Richard Freeman placed the discussion of hours worked within an international context by asking why U.S. and Canadian workers work harder (i.e., more hours per year) than their European counterparts. During the past two decades, the gap between time worked by employed North Americans and employed western Europeans increased noticeably. So too did earnings inequality. This chapter puts forth the hypothesis that higher inequality in outcomes (earnings) induces workers to work harder (i.e., longer). In Europe, where pay differences are relatively small, workers will gain little by working harder (or lose little by not working harder). In high-inequality countries, where pay differences are larger, often employment is less secure and unemployment benefits more modest; thus, there is much more of an inducement to work harder.

In support of this hypothesis, the authors observe that the greater work ethic in North America is fairly recent. From the 1950s to the 1970s, workers in the United States worked fewer hours than those in many European countries; by the 1980s and 1990s, they were working substantially longer hours than workers in all European countries. It seems unlikely that cultural differences alone could explain this change. Gross domestic product (GDP) per capita cannot explain the differences either. In some "wealthy" countries, such as Germany, employees work relatively few hours, while in others, such as Japan and the United States, they work very long hours. A corollary to this observation is that GDP per capita has some shortcoming as a measure of labor productivity and welfare. Productivity will be overestimated in countries with longer working hours, as will welfare. The additional leisure achieved by workers in countries with shorter working hours would add to a worker's utility, but this is ignored in the measure.

Turning to U.S. data on full-time workers, Bell and Freeman use a regression framework and find a clear positive association between wage inequality and hours worked. When part-time workers are included, this association is weakened. In sum, the authors argue that in the United States, and to a lesser extent in Canada, employees work hard because they face a "carrot" for doing so, and a substantial "stick" if they do not. This holds implications for labor supply models, because an inequality measure is not often included in labor supply equations.

In the research and debate on the rising earnings inequality in Canada and the United States, working time has been largely ignored. Earlier studies by Burtless (1990) and Moffitt (1990) concluded that the rise in earnings inequality in the United States is associated primarily with an increased dispersion in hourly wages, not hours of work. There appears, however, to be an association between a polarization in hours of work and that of annual earnings in Canada. In Part II, Garnett Picot reviews the evidence for this in the Canadian research literature and extends the analysis to determine whether it is changes in weeks worked per year or hours per week that explain the changing distribution of working time and its impact on earnings inequality. A series of special household surveys are used to construct comparable data on weekly hours and hourly wages over the period 1981–1993, including the Survey of Consumer Finances, the Survey of Work History, the Survey of Union Membership, the Labour Market Activity Survey, and the new Survey of Labour and Income Dynamics.

The chapter finds that the rise in inequality among prime-age males, and the decline among their female counterparts, is associated primarily with changes in hours of work, while changes in the distribution of hourly wages played a much smaller role. While changing hours plays an important role in rising inequality in general, it does not have the same effect for younger workers. Rather, declining real and relative hourly wage adjustment appears to be concentrated among the young.

Within the context of increasing wage inequality and male joblessness in the United States, Robert Haveman, Lawrence Buron, and Andrew Bershadker ask whether one can devise a more encompassing and informative measure of labor market performance than, say, the unemployment rate, or any other existing single indicator. They focus on the underutilization of male labor in the United States over the 1975–1992 period. Changes in total annual hours worked are central to their measure of underutilization. Potential earnings is defined as the amount that individuals would earn if they worked full time, all year long; this is based on a predicted hourly wage. The difference between this potential and actual earnings is defined as foregone potential earnings (FPE), the primary measure of interest in the chapter.

The FPE indicator is a more encompassing measure of changes in economic performance than the unemployment rate or changes in wages because it includes multiple dimensions of change (i.e., changes in the participation rate, the unemployment rate, hours worked, and wages paid). Furthermore, change in FPE can be decomposed into reasons, such as unemployment, illness/disability, retirement, voluntary part-time work, and housework/child care.

The authors find a declining utilization of the stock of male human capital over the 1975–1992 period. This trend was concentrated among very young workers, older workers, those with the lowest education levels, and nonwhites. In the aggregate, FPE associated with exogenous constraints (i.e., unemployment, discouraged worker effect, and illness) fell over the period, but this was more than offset by a rise in FPE due to individual responses to incentives, such as retirement, voluntary part-time work, and family responsibilities.

In a chapter concerned with spatial as well as temporal change in labor market conditions, Bob Gregory and Boyd Hunter focus on the growth of income and employment inequality in Australian cities between 1976 and 1991. Their data, which cover just over a third of the Australian population, reveal a dramatic change in that society. In an era of rising individual income inequality in Australia, inequality among neighborhoods (areas containing 200–300 dwellings) increased significantly. Among the 5 percent of neighborhoods with the lowest socioeconomic status (SES), average household income *fell* 23 percent; among the 5 percent of neighborhoods with the highest SES, income *increased* 23 percent over the period. Most of this rising spatial inequality was associated with employment changes. In 1976, employment was more or less equally distributed across neighborhoods, no matter where they fell on the SES scale. By 1991, employment in the lowest SES neighborhoods had fallen 37 percent; it was much more inequitably distributed among neighborhoods. The authors ask whether it matters if undesirable outcomes such as declining employment opportunities and falling incomes are concentrated spatially.

They suggest that greater economic polarization in our cities will increasingly lead to the emergence of "bad neighborhood" pathologies.

The chapters in Part III examine issues related to labor supply and hours constraints. To improve our understanding of labor supply changes, Richard Mueller conducts a detailed analysis of hours worked per day and days worked per week. He essentially moves beyond the traditional approach to labor supply issues, which is based on hours per week and weeks per year, and asks whether further disaggregation to hours per day and days per week makes any difference. He concludes that using weekly hours or more aggregated labor supply measures masks important differences in the labor supply decisions of individuals. His analysis shows, for example, that the reduction in female labor supply associated with having young children occurs more along the days dimension than hours per day. This suggests that the costs associated with these women supplying labor are borne more on a daily rather than an hourly basis. This would be missed in a more aggregate model. He also finds interesting associations between hours and days worked, and job change. He concludes that many job-changers desire increased flexibility. Not only do job-changers show more variability in their hours and days worked in their initial jobs, but this variability increases when they move to their new jobs.

Overall, Mueller notes that in spite of the heightened concern over "flexibility" of working time, the overwhelming norm is rigidity in weekly work schedules, particularly in days per week. He concludes that, taken together, the evidence suggests that workers desire more flexibility in their choice of hours and days worked, particularly days worked. He argues that to carry such research forward, it is necessary to disaggregate the analysis of labor supply decisions (including constraints) into its days and hours components.

Kevin Lang and Shulamit Kahn consider, both theoretically and empirically, workers' survey responses on the number of hours they would prefer to work at the same hourly rate compared with how much they actually do work. They observe that hours are constrained given the gap between preferred hours and actual hours, especially in Canada and the United States. By studying how hours constraints may determine working hours, the chapter urges caution regarding mandated hours reductions such as work sharing without first understanding why both hours constraints and unemployment exist.

The extent of hours constraints is investigated using the Panel Study of Income Dynamics, which provided the data on the relation between desired and actual hours for U.S. workers over the period 1968–1987. Results for similar surveys in Canada and the European Union are compared. To inform the empirical analysis, four primary theories in the literature are presented that may explain why workers are constrained to work more or less than they desire. These include theories on long-term contracting, hedonic wage/hours locus, fixed-wage contracts, and hours as a screening device. The analysis suggests that hours constraints are best understood in the context of the imperfect matching of wages and hours as predicted by a hedonic model. Such an imperfect matching model would allow hours policies to be evaluated in circumstances where unemployment as well as vacancies can arise. The main conclusion from their empirical analysis is that mandated hours restrictions to increase work sharing would tend to increase unemployment and lower welfare.

Part IV focuses on "short-time compensation (STC) programs." Karen Needels and Walter Nicholson study the programs' effect on how firms adjust their labor forces during cyclical downturns. Short-time compensation (sometimes called a work sharing agreement) involves providing the equivalent of unemployment insurance (UI) benefits to workers who have not been completely laid off but whose hours of work have been reduced. An evaluation of such programs needs to know what effect they have on unemployment levels, the tendency of firms to turn to hours adjustments rather than layoffs, whether they increase or decrease the total amount of compensated unemployment, and whether these programs represent an effective use of benefit payments. The authors find that neither theory nor current empirical methodologies are up to the task of providing definitive answers to such questions. Much of the theory is based on "stylized" versions of the UI and STC programs and do not allow for the complex interaction that in fact takes place between the two.

Empirical methodologies relying on self-reporting by firms of the impacts of STC often reach implausible conclusions. Methodologies that use statistical matching methods (matching similar firms in similar economic environments that do and do not use STC programs) hold more promise, as used in Canadian evaluations of work sharing. In their own study of U.S. firms based on this methodology, the authors reached two conclusions: 1) firms that use STC programs also make extensive use of layoffs, and 2) selectivity effects (i.e., unmeasured differences between firms that participate in the STC programs and those that do not) pose major, and possibly insurmountable, problems in statistical inference. Given the difficulties facing research on the effects of STC, the authors argue for more innovative use of aggregate data, random-assignment experiments, or carefully designed case studies.

To further improve the evaluation of short-time compensation initiatives, Alec Levenson draws attention to the need to account for the incidence of private short-time work (STW) arrangements in the absence of formal UI-funded work sharing programs. Based upon data drawn from the Current Population Survey (CPS), the March Annual Demographic Files, and the CPS Outgoing Rotation Group Files, his analysis shows that, compared with the very low STC take-up rates, STW was a prevalent phenomenon in the United States for the period 1968–1993. The data suggest the STW rate ranges from 50 to 70 percent of the layoff rate. The study concludes that analyses that ignore the role of STW most likely overstate the potential for layoff reductions under STC or work sharing and, as a corollary, understate the degree of public subsidy provided by such UI programs to firms that already use STW. It is recommended that existing patterns of STW be used to provide baseline estimates in future work sharing evaluations to determine the program impact on employment versus hours adjustment in the United States.

In a policy evaluation chapter on short-time compensation, Tom Siedule, Ging Wong, and Carol Guest turn to the Canadian Work Sharing Program, first implemented in 1981. They review a 1993 evaluation of the program and conclude that it clearly avoided layoffs. The evaluation suggested that about two-thirds of the layoffs that one might have expected the program to avert were in fact prevented. The program appeared to benefit both workers and firms, but it was more expensive to run (i.e., it was a bigger draw on the UI fund) than the comparable layoff alternative.

In extending this work, the authors ask whether the demand for work sharing is sensitive to changes in economic activities. If so, it is conceivable that the increased demand would create financial hardships for the UI fund during a severe economic downturn. To test this, they first estimate the likelihood of a firm participating in the program as a function of the change in the unemployment rate and of various firm characteristics. They find that work sharing is sensitive to the change (rather than the level) of economic activity and that participation is higher among firms with highly skilled workforces. Based on a macroeconomic simulation of a high-unemployment counterfactual for the 1987–1990 period, they then estimate the additional number of firms that would have used the work sharing program during this period had a more depressed economic scenario existed. They assess the additional cost of this increased demand for the program. The authors conclude that the "reasonable" depressed economic scenario chosen, which was driven by reduced exports, would have led to a 1.5-percentage-point increase in the unemployment rate by 1990. This, in turn, would have resulted in a 9 percent increase in work sharing, at a cost of $14 million. Work sharing represents such a small portion of the total UI budget (0.43 percent in 1990), such an increase (to 0.53 percent) would not be a financial burden. However, they also note that the cost of extending work sharing to all potential layoff situations could be prohibitive.

Notes

1. White, Michael. 1987. *Working Hours: Assessing the Potential for Reduction*. Geneva: International Labour Office.
2. Rifkin, Jeremy. 1995. *The End of Work: The Decline of the Global Labor Force and the Dawn of the Post-Market Era*. New York: G.P. Putnam's Sons.

Part I

1

The Changing Workweek

Trends in Weekly Hours of Work in Canada, 1976–1995

Mike Sheridan, Deborah Sunter, and Brent Diverty
Statistics Canada

There is a perception that the hours worked in the Canadian labor market have shifted away from a standard workweek. This perceived shift has been characterized as a polarization, a situation in which workers get pushed into short or long hours as the middle shrinks.

The demand for information on work hours is now greater than ever before. Economists and policymakers are interested in the relationship between the distribution of work and unemployment, particularly in light of employer reactions to legislated payroll taxes and training costs. In an economy that seems pressed to create new jobs, some observers have proposed that hours of existing jobs be redistributed to combat unemployment. Others, concerned with workplace and family stress, would like information on the joint and individual work hours and schedules of family members, who are working more weekly hours than ever before simply to maintain their standard of living. The data presented in this chapter underline the complexities involved in the development, implementation, and monitoring of policy solutions for the hours inequality and polarization phenomena.

We examine and attempt to quantify the movements away from the standard workweek by providing an analysis of Canadian trends in weekly hours of work for 1976 through 1995. Attention is devoted to 1976, 1980, 1985, 1989, and 1995 in order to eliminate, to the degree possible, discontinuities of recessions and expansions. Specifically, the analysis assesses movements in weekly hours as they relate to a number of key labor market indicators, including class of worker, age, sex, education, occupation and industry. The actual labor market factors and conditions leading to such changes are complex and remain the subject of much hypothesis, speculation, and debate.

The data presented in this chapter support the concept that, while hours worked are in flux, the distribution of hours has polarized over time. Underlying the overall trend, however, are complex and primarily unidirectional shifts in the distribution of hours, evident when workers are grouped by various characteristics. "Hours polarization" is defined here as a decline in the proportion of people working standard hours (a 35- to 40-hour week) and an increase in both the proportion working long hours (41 or more) and those working short hours (34 or less). Likewise, "hours inequality" is defined as a unidirectional shift in the distribution of hours, characterized by a decline in the proportion of people working standard hours and a corresponding increase in the proportion working *either* long or short hours (but not both).

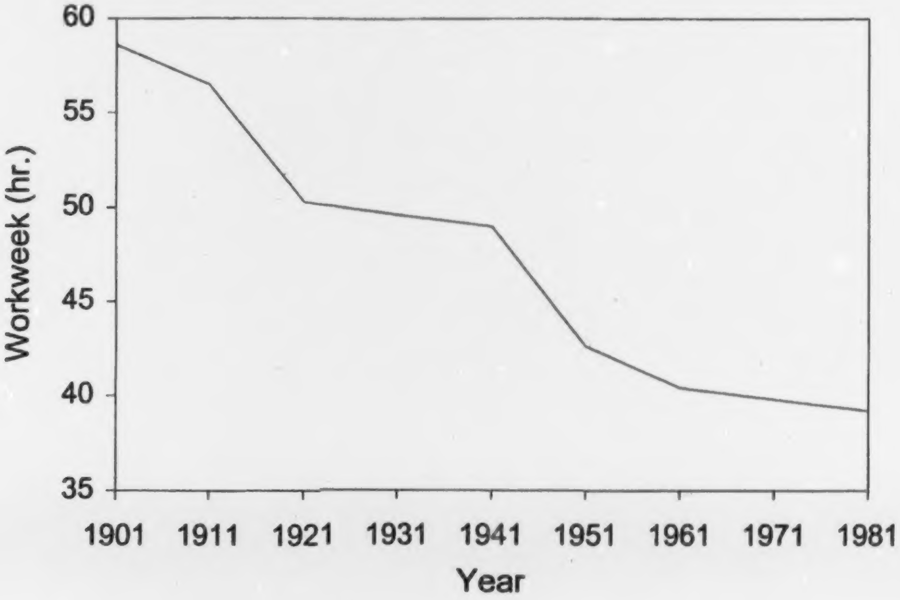
BACKGROUND—HISTORY OF HOURS WORKED IN CANADA

At the turn of the century, workers typically put in near 60-hour weeks spread over six days (Figure 1). By the 1960s, the workweek had been reduced to 37–40 hours over five days—a standard that has changed little since (Reid 1985). Increased productivity and growth in real wages spurred the trend to the shorter workweek. Workers and their employers could afford a shorter week.

The average workweek has remained fairly stable since the mid 1960s, partly because some workers have opted for nonwage benefits instead of shorter weeks. However, averages mask recent changes in the distribution of hours, especially since the 1981–1982 recession. By 1995, only 54 percent of workers in all jobs put in standard hours (35–40 hours per week), down from 65 percent in 1976. This decline occurred despite the fact that average weekly hours fell only 3 percent, from 39.0 to 37.9 hours, over the same period.

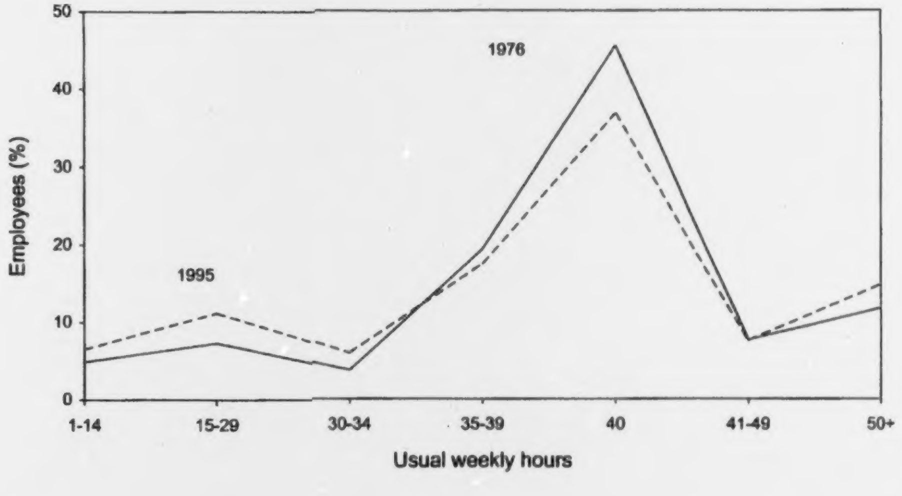
The decline in the proportion of people working standard hours was coincident with increases in the proportion of people working both long and short hours (Figure 2). As the proportion of standard hours fell (by 10.7 percentage points from 1976 to 1995), the share of workers whose usual weekly hours were less than 35 grew from 16 percent to nearly 24 percent, and that of those with 41 or more, from 19 percent

Figure 1 Standard Weekly Hours in Canada, 1901–1981



SOURCE: Reid (1985).

Figure 2 Overall Employment Hours 1976 and 1995



to 22 percent. The overall distribution of usual weekly hours polarized between 1976 and 1995. The full effect of this is not evident, however, until these movements are disaggregated by specific groups of workers. Furthermore, regardless of the changes taking place, for most industries standard hours are still the norm. What is being discussed here is changes in the poles of the hours distribution.

Men and women tend to have very different work schedules, and this is borne out in their respective hours distributions. While a roughly similar proportion of men and women worked standard hours in 1995 (55.7 percent and 52.3 percent, respectively), 2.6 times as many women than men worked short hours (35.4 percent versus 13.7 percent), while nearly 2.5 times as many men worked long hours (30.6 percent versus 12.3 percent). Despite differences in their overall distributions of hours in the poles, the two groups have faced similarly sized changes over time; that is, both groups are experiencing polarization of hours.

What has caused this apparent shift of weekly hours to the poles of the distribution? Three labor market phenomena of the 1980s and 1990s may have contributed to these changes, and to the decline in the importance of the standard workweek.

THE EFFECT OF MOONLIGHTERS, SELF-EMPLOYMENT, AND YOUTH EMPLOYMENT ON HOURS

Growth in multiple jobholding may have led to increases in the upper tail of the hours distribution, growth in self-employment could have contributed to polarization into both the upper and lower tails, and soaring school attendance rates for youths coupled with growth in part-time jobs during school could have increased the share of below-standard hours. Since these are important aspects of Canada's changing labor market, each warrants examination in any discussion of trends in work hours.

Moonlighters

Not surprisingly, moonlighting contributes to the incidence of long workweeks when hours per person are measured rather than hours per job. Since moonlighters accounted for a larger share of employment in 1995 (5 percent) than they did two decades earlier (2 percent), their tendencies to work long hours have contributed to the growth in the upper tail of the hours distribution over the period. But this influence may be on the wane. In 1995, only 64 percent of all moonlighters had above- standard workweeks, down from 70 percent in 1985 and 80 percent in 1976. The drop is explained by the fact that moonlighters in 1995 were more likely to be young persons holding down two part-time jobs that may not have added to even a standard workweek. In contrast, moonlighters 20 years ago were much more likely to work standard hours at their main jobs, so the second job was bound to push them into the long-hours category.

The Self-Employed

The second trend that may have contributed to overall hours polarization, especially to the growing share of long workweeks, is the growing prominence of self-employment, up from 12 percent of employment in 1976 to 16 percent in 1995. Moreover, its growth in the proportion of employment shows the same ratchet-like movement as overall hours polarization, with spikes during the last two recessions followed by plateaus.

The self-employed were more than twice as likely as paid employees to work long hours in 1995 (44 percent versus 15 percent). Only 3 out of 10 (compared with 6 out of 10 paid employees) worked standard hours, so their added numbers increased the upper tail of the hours distribution. On the other hand, they were also somewhat more likely to work short hours (29 percent compared with 24 percent), a tendency that has increased in the last few years. This has contributed to the lower tail of the hours distribution.

However, as with moonlighters, the self-employed have been moving away from long hours; their share of standard hours has remained stable, while that of short hours has grown considerably. The downward shift is particularly evident for those in agriculture and trade, with movement toward both standard and short hours, and for those in construction and business services, mostly into short hours.

The long-term effect of growth in self-employment on the overall distribution of hours is difficult to assess. Since much of the increase in the short-hours pole coincided with the recession of the early 1990s, that phenomenon may not persist. Furthermore, women, whose relatively small share of self-employed positions is growing, are shifting up into standard and long hours. (The overall downward trend for self-employed hours is influenced predominantly by men.)

Youths

How have youths (ages 15 to 24) affected the overall hours distribution? All movement for this group has been into the short-hours tail of the distribution. This may be largely a function of increased school attendance, since full-time students tend to work part time in order to balance work and school demands. School attendance has increased sharply since the early 1980s, rising from about 43 percent in 1984 to 57 percent in 1995.

However, short workweeks have become more common within both the student and nonstudent groups, at the expense of standard hours. In fact, 3 out of 10 employed youths who had left school worked short workweeks in 1995, triple the proportion of 1976. The change in the hours distribution for nonstudents has followed a somewhat disturbing pattern, with sharp drops triggered by the recessions and little or no gain during the recovery and expansion periods (Table 1).

The trends for these "special" workers—multiple jobholders, the self-employed, and youths—affect the overall distribution of hours and contribute to the observed polarization. However, while these groups do influence the overall trend away from standard hours, they are far from the whole story: when they are removed from the analysis, hours polarization persists. The remainder are paid employees who are at least 25 years of age and working only one job, defined here as "adult employees." This group, making up 70 percent of the total workforce in Canada in 1995, is important not only because of its size but also because of the limited scope many of its members have for controlling the hours they work.

Table 1 Distribution of Usual Weekly Hours[a] for Youths Aged 15–24 (%)

Group/weekly hours	1976	1980	1985	1989	1995	Percentage point change 1976–95
Student						
1–34	81.4	86.6	88.2	90.3	92.2	+10.8
35–40	16.4	11.6	10.3	8.6	6.7	–9.7
41+	2.2	1.8	1.5	1.1	1.1	–1.1
Nonstudent						
1–34	9.8	12.3	19.7	16.7	28.6	+18.8
35–40	76.4	74.1	66.6	68.8	57.6	–18.8
41+	13.8	13.6	13.6	14.5	13.8	0.0

[a] September to April averages.

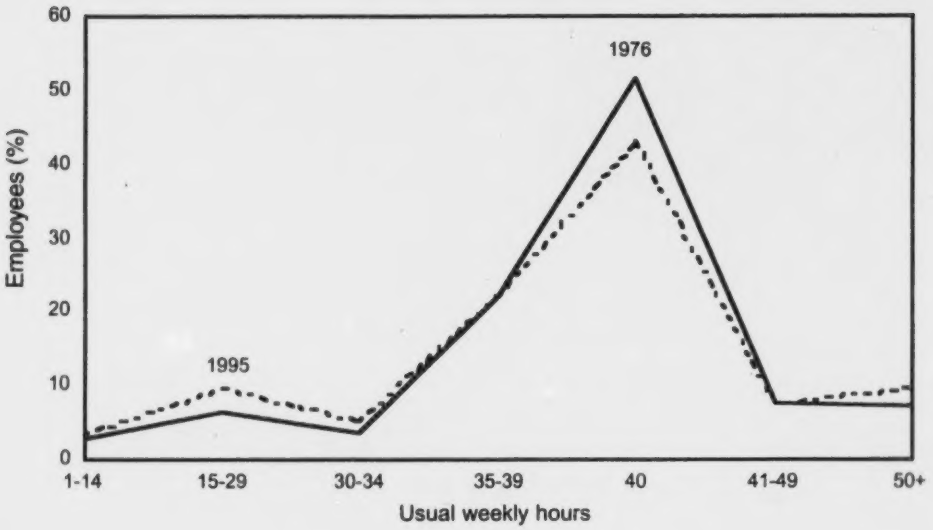
ADULT EMPLOYEES

Even after the groups who commonly work nonstandard hours are removed, polarization exists (Figure 3). In fact, these special groups account for only a small part of the decline in the number of people working standard hours. (Note that the difference between Figure 3 and Figure 2 is very small.)

A growing number of adult employees (that is, age 25 and over, not self-employed, and without a second job) are working short and long hours, with emphasis on the short hours. Of the 8.1-percentage-point decline in the proportion of adult employees working standard hours between 1976 and 1995, 5.6 points were picked up in short hours and just 2.5 points went to long. Growth in the short-hours group was steady over the period, while most of the growth in the long-hours group occurred after 1985.

Polarization has been somewhat stronger for men than for women (Table 2). The share of workers with standard hours dropped 8.5 percentage points between 1976 and 1995 (compared to 5.1 points for women), with 5.3 points moving into long hours and 3.2 points into short hours. The shift is also more skewed toward long hours for men

Figure 3 Employees Aged 25 (1976 and 1995)



than for women, whose shares of short and long hours increased in similar proportions (up 2.4 and 2.7 points, respectively). The largest part of the shift for men occurred after 1989, indicating that polarization has occurred more recently and more quickly for them than for women.

Age

Do changing demographics make a difference to the distribution of hours among adult employees? The workforce has aged over the last two decades, as "baby boomers" have moved into their forties. Theoretically, this results in a larger share of the workforce that is well-established and highly experienced than when boomers were taking their first jobs in the 1970s.

While the age pyramid may indeed be influencing the degree of overall polarization, analysis suggests that polarization is occurring for male adult workers of all age groups, while the trends are somewhat different for adult women.

Table 2 Distribution of Usual Weekly Hours,[a] by Sex (%)

Group	1976	1980	1985	1989	1995	Percentage point change 1976–95
Men						
1–34	3.9	4.4	5.2	5.2	7.1	+3.2
35–40	77.1	77.5	75.0	73.4	68.6	–8.5
41+	19.0	18.0	19.7	21.4	24.3	+5.3
Women						
1–34	27.7	29.9	30.9	29.3	30.1	+2.4
35–40	66.4	64.5	62.6	63.4	61.3	–5.1
41+	5.8	5.6	6.5	7.3	8.6	+2.7

[a] For employees 25+ years of age.

Results show that hours polarization has occurred for men in all selected age groups (Table 3). Between 1976 and 1980, the hours distributions changed very little; however, after 1980, standard hours declined steadily in each age group. The shift out of standard hours was distributed into both long and short hours, although in all cases the shift into long hours was slightly larger.

The largest shift out of standard hours was in the 55 and over group—10.9 percentage points between 1976 and 1995— although each group saw at least a 5.8-point decline in standard-hour workers. Employees in the 45–54 and the 55-and-over groups experienced the largest shifts into long hours over the period. This finding is consistent with the fact that members of these groups are most likely to be managers and/or to have seniority, both of which are increasingly associated with long hours. Growth in the proportion working short hours was largest in the 25–34 and the 55-and-over groups, where workers with short hours are more commonly found.

Unlike men, whose proportion of workers in standard hours between 1976 and 1980 changed little, women experienced significant changes during this period (Table 4). The proportion of women working a standard workweek declined for all age groups from 1976 to 1995, but the drop was greatest for those in the 25–34 and 55-and-over groups (7.1 and 11.0 percentage points, respectively). Hours for the

Table 3 Distribution of Men's Usual Weekly Hours,[a] by Age (%)

Group/weekly hours	1976	1980	1985	1989	1995	Percentage point change 1976–95
25–34 years						
1–34	3.7	4.3	5.5	5.3	8.3	+4.6
35–40	76.8	77.1	74.3	72.4	66.8	–10.1
41+	19.5	18.6	20.3	22.3	24.9	+5.4
35–44 years						
1–34	2.7	3.0	3.8	3.5	5.0	+2.4
35–40	76.4	77.3	75.8	74.7	70.6	–5.8
41+	20.9	19.6	20.4	21.8	24.4	+3.4
45–54 years						
1–34	2.0	2.9	3.8	3.3	5.2	+3.2
35–40	79.4	78.9	76.0	74.6	70.2	–9.2
41+	18.5	18.2	20.2	22.0	24.6	+6.0
55+ years						
1–34	9.1	9.2	9.7	12.1	14.2	+5.2
35–40	75.4	77.1	74.1	71.0	64.5	–10.9
41+	15.5	13.7	16.3	16.9	21.2	+5.7

[a] For employees 25+ years of age.

25–34 group have become more polarized, while for the older group short hours have become more common.

For the 35–44 and 45–54 groups, considerably less polarization has occurred, with an initial shift toward short hours between 1976 and 1980, and a movement into long hours since then. The absence of a substantial movement out of standard hours for these two groups is likely due to the large number of women moving from marginal to career jobs over this period.

Table 4 Distribution of Women's Usual Weekly Hours,[a] by Age (%)

Group/weekly hours	1976	1980	1980	1989	1995	Percentage point change 1976–95
25–34 years	23.8	25.8	26.8	25.2	27.8	+4.0
35–40	70.5	68.7	67.1	67.6	63.4	–7.1
41+	5.8	5.5	6.1	7.2	8.8	+3.0
35–44 years						
1–34	30.4	31.0	31.8	29.7	30.2	–0.2
35–40	63.8	63.3	61.1	62.7	61.6	–2.2
41+	5.8	5.7	7.2	7.6	8.1	+2.3
45–54 years						
1–34	28.8	33.0	33.8	31.0	29.3	+0.5
35–40	65.3	61.3	59.8	61.7	61.4	–3.9
41+	5.8	5.7	6.4	7.3	9.2	+3.4
55+ years						
1–34	32.1	36.1	39.0	41.4	41.8	+9.7
35–40	62.0	58.3	55.1	51.9	51.0	–11.0
41+	5.9	5.6	5.9	6.8	7.2	+1.3

[a] For employees 25+ years of age.

Education

As with most labor market outcomes, level of education plays an important role in success and, by association, exerts a very heavy influence on hours worked. Education has a strong influence on hours distribution. The higher the educational qualifications, the greater the degree of polarization; the lower the education, the higher the incidence of low hours.

It should come as little or no surprise, then, that men and women with no completed postsecondary school qualifications have experienced the greatest increase in the short-hours tail of the distribution. The incidence of short workweeks has traditionally been relatively high for women without a completed formal postsecondary education, and that relationship has intensified (Table 5). The 5.2-percentage-point decline in the proportion of women working standard hours moved almost entirely to short hours, pushing the share of the latter up from 28 percent in 1976 to 32 percent in 1995. Men in the same educational group saw even greater losses in the 35–40 hour category over the period, with a decline of 9 percentage points in these standard hours (Table 6). However unlike those of their female counterparts, men's losses have been distributed equally between long and short hours (both up 4.6 percentage points).

At the other end of the formal education spectrum, university graduates have experienced a unidirectional shift toward long hours since 1976. Women with degrees have seen virtually no change in short hours, and men have seen only a very slight increase. Between 1976 and 1995, women in this group added about 5 percentage points to the number of long hours they worked, while men increased theirs by more than 6 full points, from 25.3 percent to 31.8 percent. It seems that long hours are part of the baggage of a higher education, especially for men.

INDUSTRY

Looking at the hours distribution across industries is another way to help shed some light on both the extent and complexity of polarization and inequality in the workplace. It may be that changes in the distribution of hours reflect structural change, with disproportional growth in industries that tend to use part-timers and in those that require long hours. Alternatively, some industries may be making increased use of short-hours workers only, while others are becoming more reliant on long workweeks. Finally, changes in the hours distribution may be spread fairly evenly within all or most industries, suggesting pervasive and systemic factors that are economy-wide.

For men working in the goods-producing sector, the shift is toward long hours (Table 7; for detailed industry data, see Appendix Table A1). Of the 10.1-percentage-point shift out of standard hours over the 1976–1995 period, 8.3 points moved to long hours while only 1.8 moved to short hours. The proportion of standard-hours workers fell in

Table 5 Distribution of Women's Usual Weekly Hours,[a] by Education (%)

Highest level of education	1976	1980	1985	1989	1995	Percentage point change 1976–95
Less than postsecondary certificate, diploma, or degree						
1–34	28.0	30.4	31.6	30.9	32.0	4.0
35–40	66.8	64.8	63.6	63.7	61.6	–5.2
41+	5.1	4.8	4.9	5.5	6.3	1.2
Postsecondary certificate/diploma						
1–34	28.7	31.5	33.5	29.8	31.2	2.5
35–40	66.3	64.1	61.1	64.2	62.6	–3.7%
41+	5.0	4.4	5.5	6.0	6.1	1.1
University degree						
1–34	23.7	24.8	24.9	22.6	23.9	0.2
35–40	63.7	62.9	60.3	61.6	58.3%	–5.4
41+	12.7	12.3	14.8	15.8	17.9	5.2

[a] For employees 25+ years of age.

Table 6 Distribution of Men's Usual Weekly Hours,[a] by Education (%)

Highest level of education	1976	1980	1985	1989	1995	Percentage point change 1976–95
Less than postsecondary certificate, diploma, or degree						
1–34	3.4	3.8	5.2	5.2	8.0	4.6
35–40	78.2	79.0	76.5	75.1	69.0	–9.2
41+	18.4	17.2	18.3	19.8	23.0	4.6
Postsecondary certificate/diploma						
1–34	3.8	4.3	4.7	4.1	5.9	2.1
35–40	80.7	80.3	78.2	76.6	72.8	–7.9
41+	15.5	15.4	17.0	19.3	21.4	5.8
University degree						
1–34	6.8	7.4	5.8	6.2	7.2	0.4
35–40	67.9	68.8	66.7	64.8	61.0	–6.9
41+	25.3	23.8	27.4	28.9	31.8	6.5

[a] For employees 25+ years of age.

Table 7 Distribution of Men's Usual Weekly Hours,[a] by Industry (%)

Industry	1976	1980	1985	1989	1995	Percentage point change 1976–95
Goods-producing						
1–34	1.5	1.9	2.4	2.3	3.3	+1.8
35–40	82.2	82.2	79.9	78.3	72.1	–10.1
41+	16.3	15.8	17.7	19.4	24.6	+8.3
Service-producing						
1–34	5.8	6.3	7.2	7.2	9.5	+3.7
35–40	73.1	74.0	71.7	69.9	66.4	–6.7
41+	21.1	19.7	21.1	22.9	24.1	+3.0

[a] For employees 25+ years of age.

every goods-producing industry over the period, with the exception of agriculture. Primary industries (mining, forestry, and fishing) in particular have seen a dramatic increase in the proportion of long-hours workers (up 19.5 percentage points between 1985 and 1995). Manufacturing and utilities have also been using proportionately more long-hours workers. Hours have been polarizing in construction and have been shifting from long to standard in agriculture.

Within the service-producing industries, the shift for men has been into both short and long hours. Polarization has been greater here than in the goods-producing sector. The 6.7-percentage point decrease in the proportion of standard-hours workers since 1976 has been evenly divided between short and long hours. Hours in transportation, storage, and communication; trade; and business services have been polarizing. The share of short hours in health and social services; accommodation, food, and beverage services; and other services has grown as has that of long hours in finance, insurance, and real estate (FIRE); educational services; and government services.

As with men, the hours distribution for women working in goods-producing industries has shifted toward long-hours workers (Table 8; for detailed industry data, see Appendix Table A2). Overall, the proportion of women working standard hours declined 4.9 percentage points between 1976 and 1995, 4.0 points of which went to long hours. The increase in the proportion of long-hours workers occurred in all goods-producing industries with the exception of agriculture.

Table 8 Distribution of Women's Usual Weekly Hours,[a] by Industry (%)

Industry	1976	1980	1985	1989	1995	Percentage point change 1976–95
Goods-producing						
1–34	13.7	15.0	16.1	15.8	14.5	+0.9
35–40	80.5	78.8	77.5	77.1	75.6	–4.9
41+	5.8	6.2	6.4	7.0	9.9	+4.0
Service-producing						
1–34	31.1	33.4	33.8	31.9	32.8	+1.7
35–40	63.1	61.2	59.7	60.7	58.8	–4.3
41+	5.8	5.5	6.5	7.4	8.3	+2.5

[a] For employees 25+ years of age.

Service-producing industries have experienced polarization, with the 4.3-percentage-point decline in the proportion of standard-hours workers being picked up by short and long hours (1.7- and 2.5-point increases, respectively). Despite an apparent overall polarization, no single industry in this group has increased its proportion of both short and long hours. Polarization is the net effect of women tending toward long hours in transportation, storage, and communication; trade; FIRE; business services; educational services; and government and other services, plus a shift toward short hours in health and social services and accommodation, food and beverage services.

The decline in standard weekly hours holds generally across all industries for both sexes, but increases in short and long hours are not always of a similar magnitude: polarization is not widespread within industries. Thus, the overall observed polarization masks underlying unidirectional changes in the hours distribution within industries, with some tending exclusively toward long hours, and others exclusively toward short hours. One clear trend for both men and women is a shift toward more long hours in goods-producing industries and a polarization in service-producing industries.

OCCUPATIONS

The number of hours worked is heavily dependent on the type of work performed. Variation in hours worked across occupations may be caused by many factors: the level of responsibility of the position, the skill level required, the cost of training new employees, the opportunity for paid overtime, and the prospects of promotion.

For men, the proportion of employees working standard hours has been shrinking in all 10 occupational groups (Table 9). Instead of widespread polarization, however, a shift in one direction is more likely, depending on occupation. White-collar[1] and blue-collar[2] occupations, for example, have experienced growth mainly in the proportion of workers with long hours. The largest shift out of standard hours between 1985 and 1995 occurred in the managerial category, an 8.2-percentage-point decline in standard-hours workers and a corresponding 8.1-percentage-point increase in the proportion of long-hours

Table 9 Distribution of Men's Usual Weekly Hours,[a] by Occupation (%)

Occupation/ weekly hours	1976	1980	1985	1989	1995	Percentage point change[b] 1985–95
Managerial						
1–34	3.2	3.1	2.6	2.4	2.7	0.1
35–40	77.0	75.2	68.3	64.9	60.0	–8.3
41+	19.7	21.7	29.1	32.7	37.3	+8.2
Professional						
1–34	7.5	8.1	7.0	8.0	9.6	+2.6
35–40	73.5	74.0	73.0	70.8	67.2	–5.8
41+	19.0	17.9	20.0	21.2	23.2	+3.2
Clerical						
1–34	4.7	5.2	5.9	6.3	9.9	+4.1
35–40	88.2	87.7	86.9	85.6	80.7	–6.1
41+	7.0	7.1	7.2	8.1	9.3	+2.1
Sales						
1–34	3.6	5.3	7.7	5.7	9.7	+2.0
35–40	63.9	65.8	63.5	64.2	62.2	–1.3
41+	32.5	28.9	28.9	30.2	28.1	–0.8
Service						
1–34	7.3	8.3	12.0	11.7	17.2	+5.1
35–40	72.1	73.1	71.5	70.8	66.3	–5.2
41+	20.7	18.7	16.5	17.5	16.5	0.0
Primary occupations						
1–34	4.4	4.0	5.5	4.9	6.9	+1.4
35–40	60.9	64.5	60.3	60.7	53.9	–6.4
41+	34.7	31.6	34.3	34.4	39.3	+5.0

(continued)

Table 9 (continued)

Occupation/ weekly hours	1976	1980	1985	1989	1995	Percentage point change[b] 1985–95
Processing, machining & fabricating						
1–34	1.0	1.4	1.8	1.9	2.3	+0.5
35–40	84.8	85.5	84.9	83.6	79.2	–5.7
41+	14.2	13.1	13.3	14.5	18.5	+5.2
Construction trades						
1–34	2.0	2.0	3.6	3.0	5.4	+1.8
35–40	83.6	85.2	82.9	78.9	74.9	–8.0
41+	14.4	12.7	13.6	18.1	19.7	+6.1
Transport operator						
1–34	5.8	6.2	8.2	8.4	9.6	+1.5
35–40	63.6	64.4	59.5	58.4	52.4	–7.1
41+	30.6	29.5	32.4	33.2	38.0	+5.7
Material handling & other crafts						
1–34	2.3	3.0	5.4	6.0	7.7	+2.3
35–40	86.1	86.1	83.5	80.8	78.0	–5.5
41+	11.6	11.0	11.1	13.2	14.4	+3.2

[a] For employees 25+ years of age.

[b] The 1984 reclassification of SOC codes included a new definition of managers, which meant that more people were classified as such. As a consequence, meaningful comparisons can be made only as far back as 1985.

workers. Clerical, sales, and service jobs, on the other hand, experienced growth in the proportion of short-hours workers, with little or no growth in the long-hours tail.

Those occupations in which long hours have become more common for men have either a high level of responsibility (white-collar jobs), or regular opportunities to work paid overtime (blue-collar). In the case of blue-collar occupations, given administrative and overhead considerations, it may be more cost-efficient for employers to pay overtime wages than to hire and train new employees. White-collar workers, especially managers, may be working longer hours because of increased responsibilities in the wake of corporate downsizing, or simply to keep their jobs in an increasingly competitive employment market. By contrast, those occupations in which short hours have become more common (clerical, sales, and service) are often low-paying and/or part-time.

Unidirectional shifts, as opposed to polarization, have also taken place in women's distribution of hours (Table 10). These shifts are generally not as strong as those for men. The proportion of standard-hours workers has declined, however, in six out of eight occupational groups.[3] Similar to the situation for men, female white-collar managers and professionals have seen growth in the long-hours tail of the distribution, while blue-collar, clerical, sales, and service occupations have exhibited no distinct pattern.

Of all occupational groups considered here, both male and female managers experienced the largest growth in the long-hours tail. It is not surprising to find that managers work long hours, nor that weekly hours increase with the level of management. Indeed, according to recent data from the Survey of Labour and Income Dynamics (SLID), nonmanagerial employees averaged 36 hours per week, while lower managers averaged 39, middle managers 40, and senior managers 42. This pattern has become more marked over time. Managers' expanding work hours, and their growing numbers in the labor market—16.7 percent of adult employees identified themselves as managers in 1995, up from 14.1 percent 10 years earlier—may be a driving force behind the overall movement into long workweeks. To determine to what extent this is the case, the hours distribution of nonmanagerial adult employees is examined here.

Table 10 Distribution of Women's Usual Weekly Hours,[a] by Occupation (%)

Occupation/ weekly hours	1976	1980	1985	1989	1995	Percentage point change[b] 1985–95
Managerial						
1–34	12.6	11.6	12.0	11.5	12.2	+0.2
35–40	81.4	80.9	76.1	75.2	72.8	–3.3
41+	6.1	7.5	11.8	13.3	15.0	+3.2
Professional						
1–34	28.2	31.8	32.8	32.8	33.2	+0.4
35–40	64.0	61.1	57.9	57.3	55.8	–2.1
41+	7.8	7.1	9.3	9.9	11.1	+1.7
Clerical						
1–34	25.2	27.8	28.3	28.2	30.0	+1.7
35–40	72.6	69.9	69.6	69.2	67.1	–2.5
41+	2.2	2.3	2.1	2.7	2.9	+0.7
Sales						
1–34	44.6	43.4	49.6	41.2	41.6	–8.0
35–40	47.3	48.7	42.2	49.0	50.7	+8.5
41+	8.1	7.9	8.2	9.8	7.6	–0.5
Service						
1–34	41.0	44.3	49.0	45.8	48.8	–0.2
35–40	50.5	48.5	44.7	47.4	43.9	–0.8
41+	8.5	7.2	6.3	6.8	7.4	+1.0
Primary occupations						
1–34	42.7	46.9	40.7	41.7	34.8	–5.9
35–40	32.3	31.8	35.8	36.5	44.1	+8.3
41+	25.0	21.3	23.5	21.8	21.1	–2.4

Occupation/ weekly hours	1976	1980	1985	1989	1995	Percentage point change[b] 1985–95
Processing, machining & fabricating						
1–34	7.4	8.3	9.6	9.6	9.2	–0.4
35–40	86.7	85.1	84.3	84.4	81.9	–2.4
41+	5.9	6.6	6.1	5.9	8.9	+2.8
Material handling & other crafts						
1–34	13.6	16.5	20.0	18.6	23.6	+3.6
35–40	82.8	79.5	76.0	76.2	69.5	–6.5
41+	3.6	4.0	4.0	5.2	6.9	+3.0

[a] For employees 25+ years of age.

[b] The 1984 reclassification of SOC codes included a new definition of managers, which meant that more people were classified as such. As a consequence, meaningful comparisons can be made only as far back as 1985.

Even after the managerial group has been removed from the analysis, a small amount of polarization remains (Figure 4).[4] The decline in the proportion of standard-hours workers is 4.5 percentage points, with 2.1 points going to short and 2.3 to long. Remaining increases in the proportion of long-hours workers are for the most part in the goods-producing industries, especially nonagricultural primary. That the shift into long hours in these industries persists even after managers have been removed indicates that overtime work by blue-collar workers is a contributing factor. Conversely, since increases in the proportion of short-hours workers are generally in the service-producing industries, with no corresponding increase in long hours, it may be concluded that growth into long hours in these industries was exclusive to managers. Polarization still exists for men, while it is virtually non-existent for women (Table 11).

Figure 4 Work Hours for Nonmanagerial Employees Aged 25+ (1976 and 1995)

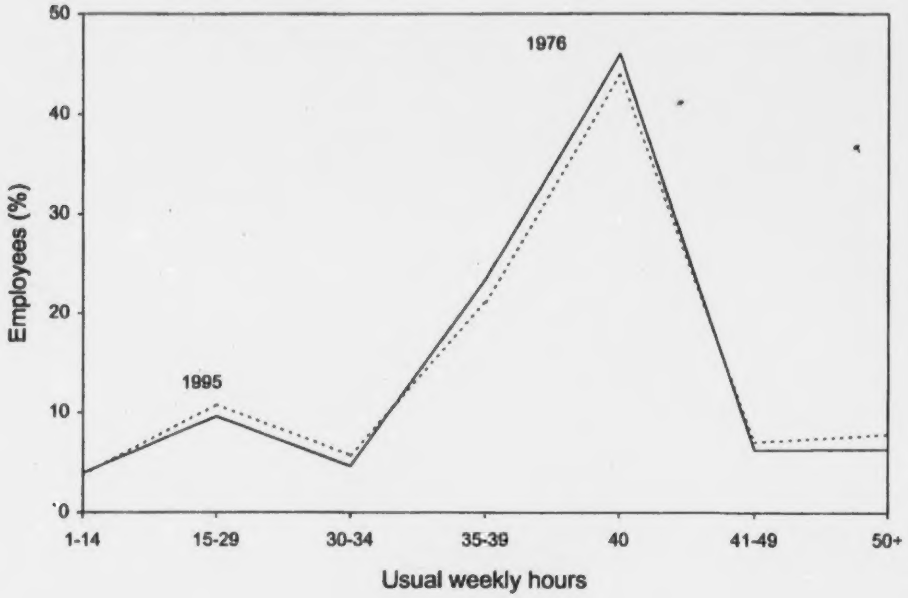


Table 11 Distribution of Usual Weekly Hours for Nonmanagerial Employees[a](%)

Group/weekly hours	1976	1980	1985	1989	1995	Percentage point change[b] 1985–95
Men						
1–34	4.0	4.6	5.8	5.7	8.0	+2.2
35–40	77.1	77.9	76.4	75.0	70.4	–6.0
41+	18.9	17.5	17.9	19.3	21.6	+3.7
Women						
1–34	28.5	31.2	33.2	31.8	33.6	+0.4
35–40	65.7	63.4	61.0	61.7	59.1	–1.9
41+	5.8	5.5	5.8	6.4	7.3	+1.5

[a] For employees 25+ years of age.

[b] The 1984 reclassification of SOC codes included a new definition of managers, which meant that more people were classified as such. As a consequence, meaningful comparisons can be made only as far back as 1985.

SUMMARY AND CONCLUSIONS

Are there major changes in the work hours of Canadians? Specifically, is the work "pie" becoming more unevenly divided between those with short workweeks and those with long hours?

Evidence from the last 20 years suggests that hours have shifted from standard to both long and short workweeks, especially since the early 1980s. This phenomenon persists, though to a lesser extent, when special groups such as the self-employed, moonlighters, and young workers, as well as managers, are removed from the analysis. The picture is not an even one. Workweeks seem to have polarized for both sexes, though somewhat more so for men, as women increasingly opt for standard or long workweeks. Evidence by age group suggests that polarization is widespread, although the shift from standard to long hours is more marked for those aged 35 to 54 than for those in the younger or older age groups. This is especially so for women.

In terms of timing, the shift from standard to short workweeks appears to have been triggered by the economic downturns of both the early 1980s and 1990s, while most of the growth in the share of long workweeks appears to be primarily a recent phenomenon.

Not surprisingly, education is strongly related to the length of the workweek, with long hours a frequent occurrence among those with higher credentials. Trends in the distribution of hours have also differed by educational level. Both male and female university graduates are increasingly likely to work long hours at the expense of standard hours. Women without any postsecondary qualifications are increasingly likely to work below standard hours, while their male counterparts are experiencing a growth in both short and long hours.

Change in the distribution of work hours is not simply a structural phenomenon. There have been measurable shifts out of standard workweeks in most industries, although the result is not always polarization. As one might expect, almost all service-producing industries have tended toward a shorter workweek, while many goods-producing industries, particularly manufacturing and other primary, have tended to long workweeks. Only a few demonstrate a clear trend to polarization: construction; transportation, storage and communication; trade; and business services.

Occupation plays a strong role in determining a weekly work schedule. There has been a marked trend toward long hours for managers, especially during the 1990s. But they are not the only ones for whom the week has grown longer. Factory workers and those in primary occupations are also increasingly likely to work more than 40 hours a week. In contrast, hours for those in sales and services are increasingly likely to be below standard.

The link between managers and the trend toward long hours has warranted closer examination. One distinguishing feature of a manager's job is supervisory responsibility, which means that hours of work will likely be above standard. In fact, it appears that the workweek lengthens as the number of persons requiring supervision increases.

Since managers are a growth occupation within almost all industries, their removal from the analysis of industry trends helps illuminate how much influence they have on overall patterns. Interestingly, when managers are removed, the direction of shifts in the hours distribution holds for most industries, although the magnitude toward long hours tends to be somewhat less. In contrast, the shift to short hours tends to strengthen in their absence.

The shift from standard hours continues to have deep and lasting implications for employers, workers, and the unemployed. Why are weekly hours becoming more unevenly distributed? While this chapter has not revealed any one causal factor, the findings are at least consistent with a number of popular hypotheses. First, the data support the contention that many employers in a variety of industries are relying more on a core group of highly educated, experienced workers—primarily managers but also those skilled in trades. Expectations for performance may be increasing in the difficult labor market of the 1990s, and core workers may simply be putting in extra hours on a regular basis to stay afloat, or as an investment in future reward through the internal promotion ladder. Second, hours are heading down in a number of industries. Most are distinguished by a requirement for relatively unskilled workers for whom job-specific training can be minimal. In this situation, workers may be treated as roughly interchangeable.

Perhaps the most important issue to emerge from the hours polarization/inequality debate is the question of the potential to redistribute hours in the labor market. One of the key questions in this debate has been, and will probably continue to be, how many of these long hours could be redistributed to those who are currently underemployed or unemployed? In the "tough labor market" conditions of the 1990s, do employees who work long hours get overtime compensation, or are they in fact doing more for the same paycheck in order to keep a job? If so, can unpaid hours be redistributed?

The other key question in this debate concerns workers at the other pole of the hours distribution. Do short-hours workers possess the skill mix and portability to assume jobs normally associated with longer hours, and should this transfer of work be achieved by squeezing the long-hours side of the pole? Without some sort of restructuring of hours, these workers may become stuck in low-end, poor-paying jobs. Further, these jobs may continue to move further in the direction of the short-hours pole.

DATA SOURCES AND DEFINITIONS

Data in this chapter are derived from two sources: the majority are annual averages derived from the monthly Labour Force Survey. Data on hours worked by supervisory and management responsibilities have been drawn from the 1993 Survey of Labour and Income Dynamics.

Usual hours: the number of hours usually worked by a respondent in a typical week, regardless of the nominal schedule and regardless of whether or not the hours are paid.

Polarization: a decline in the share of standard hours (35 to 40 per week) with roughly equal gains in the share of short (less than 35 hours) and long (41 and over) workweeks.

Standard hours: 35 to 40 hours per week. Coincides with a notion of adequate employment: lower than legislated thresholds for overtime pay, but high enough to assure eligibility for benefits.

Self-employed: includes all working owners, whether or not they are incorporated or have paid help. Also included are family members who work for a family business without pay.

Inequality: refers to a unidirectional shift in the distribution of hours from standard to long or short.

Notes

1. Managers and professionals.
2. Nonmanagerial employees in primary occupations, processing, machining, fabrication, construction trades, transport operation, and materials handling.
3. Occupational groups such as construction trades and transport operations were dropped due to small sample sizes.
4. The amount of polarization is likely to be understated somewhat because it is being measured from 1985 instead of 1976. As mentioned previously, a change in the definition of manager in 1984 has created a break in the data series by occupation.

Reference

Reid, F. 1985. "Reductions in Work Time: An Assessment of Employment Sharing to Reduce Unemployment." In *Work and Pay: The Canadian Labour Market*, W.C. Riddell, ed. Collected Research Studies/Royal Commission on the Economic Union and Development Prospects for Canada, 17. Toronto: University of Toronto Press, pp. 141–169.

Table A1 Distribution of Men's Weekly Hours Worked,[a] by Industry (%)

Industry/weekly hours	1976	1980	1985	1989	1995	Percentage point change 1976–95
Agriculture						
1–34	9.5	12.0	10.4	12.4	11.8	+2.2
35–40	28.6	29.4	33.1	32.1	40.3	+11.8
41+	61.9	58.6	56.5	55.5	47.9	–14.0
Other primary						
1–34	1.0	1.1	1.7	1.5	2.0	+1.1
35–40	79.6	77.9	73.4	71.7	58.9	–20.6
41+	19.5	21.0	24.9	26.9	39.0	+19.5
Manufacturing						
1–34	1.1	1.5	1.8	1.8	2.3	+1.2
35–40	85.5	84.9	83.4	81.9	76.1	–9.4
41+	13.4	13.5	14.8	16.3	21.5	+8.2
Construction						
1–34	1.9	2.8	4.6	3.9	6.6	+4.7
35–40	76.6	77.6	72.2	71.3	65.2	–11.4
41+	21.5	19.6	23.2	24.9	28.2	+6.7
Utilities						
1–34	1.0	1.5	1.0	1.6	1.8	+0.8
35–40	93.9	92.4	93.2	90.5	85.8	–8.1
41+	5.2	6.1	5.8	7.9	12.4	+7.3
Transportation, storage, and communication						
1–34	4.6	4.2	5.6	6.5	7.3	+2.8
35–40	78.1	78.0	75.2	71.7	66.4	–11.7
41+	17.3	17.8	19.2	21.9	26.2	+9.0
Trade						
1–34	3.1	3.5	5.5	4.5	7.4	+4.3

(continued)

Table A1 (continued)

Industry/weekly hours	1976	1980	1985	1989	1995	Percentage point change 1976–95
35–40	68.4	71.0	69.5	68.9	64.7	–3.7
41+	28.5	25.4	25.0	26.6	27.9	–0.6
Finance, insurance, and real estate						
1–34	6.2	7.2	6.8	5.9	6.7	+0.4
35–40	69.6	69.1	65.3	64.5	65.1	–4.5
41+	24.2	23.7	27.8	29.6	28.2	+4.0
Business services						
1–34	5.4	4.9	6.5	5.2	7.8	+2.5
35–40	74.7	74.5	70.0	69.3	66.2	–8.4
41+	20.0	20.6	23.5	25.5	25.9	+6.0
Educational services						
1–34	11.4	11.9	10.8	11.4	13.5	+2.1
35–40	66.9	67.8	65.1	63.2	59.4	–7.5
41+	21.7	20.4	24.1	25.4	27.2	+5.5
Health and social services						
1–34	7.0	8.4	10.3	12.1	14.5	+7.5
35–40	67.9	70.8	70.6	69.5	68.1	+0.3
41+	25.1	20.8	19.1	18.4	17.3	–7.8
Accommodation, food, and beverage services						
1–34	8.1	11.1	15.0	15.2	21.8	+13.7
35–40	55.1	58.2	56.7	55.9	54.6	–0.5
41+	36.8	30.7	28.3	28.9	23.6	–13.2

Industry/weekly hours	1976	1980	1985	1989	1995	Percentage point change 1976–95
Other services						
1–34	10.1	11.9	14.6	14.7	17.2	+7.1
35–40	64.6	66.0	60.4	56.9	59.5	–5.1
41+	25.4	22.0	25.0	28.4	23.3	–2.0
Government services						
1–34	5.0	5.4	4.3	4.3	4.9	–0.1
35–40	86.0	85.7	85.2	84.5	81.7	–4.3
41+	9.0	8.9	10.5	11.2	13.4	+4.4

[a] For employees 25+ years of age.

Table A2 Distribution of Women's Weekly Hours Worked,[a] by Industry (%)

Industry/weekly hours	1976	1980	1985	1989	1995	Percentage point change 1976–95
Agriculture						
1–34	47.4	50.8	48.2	49.1	40.8	–6.7
35–40	30.3	26.9	31.9	34.2	41.7	+11.4
41+	22.3	22.3	19.8	16.8	17.6	–4.8
Other primary						
1–34	19.3	15.0	12.1	18.1	17.3	–2.0
35–40	75.4	79.8	80.9	70.1	67.7	–7.7
41+	5.2	5.2	7.0	11.7	15.0	+9.7
Manufacturing						
1–34	9.6	10.3	10.9	10.8	10.5	+0.9
35–40	85.6	84.3	83.7	82.8	80.1	–5.4
41+	4.9	5.4	5.4	6.4	9.4	+4.5
Construction						
1–34	34.6	45.1	45.4	38.9	35.2	+0.6
35–40	60.3	49.3	47.7	55.5	56.8	–3.5
41+	5.1	5.6	6.9	5.6	8.0	+3.0
Utilities						
1–34	11.3	5.0	8.9	7.5	9.2	–2.1
35–40	87.1	94.4	90.7	90.6	84.8	–2.3
41+	1.7	0.6	0.4	2.0	6.1	+4.4%
Transportation, storage, and communication						
1–34	28.0	28.8	25.5	23.9	25.0	–3.0
35–40	68.6	66.7	70.3	71.3	68.2	–0.4
41+	3.4	4.4	4.2	4.9	6.9	+3.5

Industry/weekly hours	1976	1980	1985	1989	1995	Percentage point change 1976–95
Trade						
1–34	40.7	43.5	43.6	39.0	40.3	–0.5
35–40	54.1	51.7	51.0	54.4	52.3	–1.9
41+	5.2	4.8	5.4	6.5	7.5	+2.3
Finance, insurance, and real estate						
1–34	21.6	19.7	18.8	19.7	20.6	–0.9
35–40	73.2	75.2	74.5	73.0	72.0	–1.2
41+	5.2	5.2	6.7	7.3	7.3	+2.1
Business services						
1–34	26.9	26.7	25.1	23.4	22.1	–4.8
35–40	70.8	68.7	68.5	69.0	68.5	–2.3
41+	2.2	4.6	6.4	7.6	9.4	+7.1
Educational services						
1–34	32.0	35.1	32.3	31.9	32.3	+0.2
35–40	56.8	54.8	54.0	53.4	50.1	–6.7
41+	11.1	10.1	13.7	14.7	17.6	+6.5
Health and social services						
1–34	26.2	33.1	38.3	38.6	39.4	+13.2
35–40	70.3	63.6	58.3	57.3	56.2	–14.1
41+	3.5	3.2	3.4	4.0	4.5	+0.9
Accommodation, food, and beverage services						
1–34	39.8	42.6	48.2	41.9	46.1	+6.2
35–40	49.6	48.8	43.2	48.1	44.7	–4.9
41+	10.6	8.6	8.6	10.0	9.2	–1.3
Other services						
1–34	46.8	44.1	41.6	37.5	38.2	–8.6
35–40	45.8	48.1	47.5	50.9	50.3	+4.5
41+	7.5	7.8	10.9	11.6	11.5	+4.1

(continued)

Table A2 (continued)

Industry/weekly hours	1976	1980	1985	1989	1995	Percentage point change 1976–95
Government services						
1–34	14.7	15.2	16.8	15.2	15.0	+0.3
35–40	83.5	82.7	80.6	81.0	80.7	–2.8
41+	1.8	2.1	2.7	3.8	4.3	+2.5

[a] For employees 25+ years of age.

2

Trends in Hours of Work in the United States

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Efforts to shorten and standardize the length of the workweek were at the forefront of labor market issues in the first four decades of this century, culminating in the enactment of the Fair Labor Standards Act of 1938. After long and hard-fought legal and political battles, the act allowed for a maximum workweek of 44 hours, to decline to 40 hours in the third year after enactment (Elder and Miller 1979). Although employers could still demand longer workweeks, hours worked beyond the legal maximum would require time-and-a-half pay.

While workweek issues have fallen from the fore in recent decades, they still touch many key labor market topics and trends. For example, arguably the two most dominant trends in the post–World War II work world have been the influx of women, particularly mothers, into the job market, and the steady decline in the retirement age. Women have increased their numbers in the workforce, and they have also shifted their work schedules toward year-round, full-time employment. In addition, as work activity among older men was declining, those left working were increasingly likely to work part time.

The key labor issues of the day in the 1990s were most likely worker displacement and the quality of jobs. These issues, too, have workweek components. Even as the overall U.S. employment numbers have risen substantially, millions of jobs have been lost each year to corporate and government restructuring. A common perception is that those spared such job loss, particularly those in managerial and professional jobs, have been compelled to work even harder—longer—to protect their jobs. As for the quality of jobs, new jobs created often are stereotyped as part-time, low-wage, poor-quality jobs.[1]

This chapter examines work-hour trends from two perspectives. First, trends in the average workweek and changes in the distribution of hours worked since the mid 1970s are examined. Then, the focus is expanded to estimate annual work hours. This figure is affected by the extent to which people work at all and the number of weeks they work during the year, in addition to the length of the workweek. Lastly, the appendix provides a discussion of the differences in hours data collected following the redesign of the Current Population Survey (CPS), which was implemented in January 1994, from those obtained prior to 1994. Because of the effect of those changes on work-hour estimates, trend data in the chapter are restricted to the period through 1993.[2]

MEASURING HOURS OF WORK

Estimates of the length of the workweek can be obtained from workers themselves or from their employers. Employer-based surveys count the total number of jobs held by workers, so average hours calculated from those data are reported per job, not per worker. Workers, of course, can work at more than one job. Also, workweek estimates from employers generally are for hours paid (including paid annual and sick leave) rather than hours actually worked. Another shortfall of employer-based surveys for this analysis is that they typically lack demographic information—such as age, gender, and education—that are critical to understanding workweek trends. Thus, if the focus is on workers and their work schedules, employer surveys will not suffice.[3]

For those reasons, data obtained from individuals will be used in this analysis. The CPS provides comprehensive and consistent hours-at-work and employment time-series data that can be obtained for many demographic characteristics.[4] Respondents to the survey are queried on their usual and actual hours at work. Additionally, respondents surveyed in March are asked about their work experiences in the prior year, including their typical work schedules and the number of weeks worked. The analysis generally is limited to nonagricultural wage and salary workers.[5]

AVERAGE HOURS AT WORK

In 1995, the average workweek for nonagricultural wage and salary workers was 39.2 hours. That average varies considerably across worker groups, however. For instance, the average workweek for men was 42.1 hours compared to 35.8 hours for women, and persons aged 25 to 54 typically work longer than do younger and older workers (Table 1). In addition, the length of the workweek varies by marital status. Married men have the longest workweek and, in 1995, worked an average of eight hours per week more than married women. Reflecting their younger age, both men and women who were never married worked the shortest workweek.[6]

Average hours at work changed little over the period from 1976 to 1993, only increasing by 1.1 hours, on net, to 39.2 hours.[7] But during this period, the age distribution of the U.S. working population changed substantially and in a way that influenced the length of the average workweek. By 1993, the baby boomers—those born between 1946 and 1964—all had moved into the central working ages of 25 to 54. Meanwhile, workers in the younger and older age groups, which include many students and retirees, comprised a declining share of employment. Workweeks typically are the longest for workers aged 25 to 54, and part-time (and part-year) employment is most common among younger and older workers. These shifts in the age distribution, then, would tend to increase the length of the average workweek, all other things being equal.

To determine the effect of the shifting age distribution of the employed on the change in the average workweek for men and women, it is necessary to calculate average hours in 1993 assuming that the age distribution of those at work had remained unchanged since 1976.[8] As Table 2 shows, after removing the effect of the shifting age distribution, men had virtually no rise in their average weekly hours (edging up from 41.0 to 41.2 hours), and women's average workweeks rose by only an hour.

The small changes in the length of the workweek, whether on an age-adjusted or unadjusted basis, reflect (and mask) offsetting increases and decreases in the hours-at-work distribution. As shown in Figure 1, between 1976 and 1993, the proportion of nonagricultural

Table 1 Nonagricultural Wage and Salary Workers at Work and Their Average Hours, by Age, Sex, Race, and Hispanic Origin, 1995 Averages

Characteristic	Total at work (000)	Average hours: Total at work	Average hours: Persons who usually work full time
Age and sex			
Total	107,656	39.2	43.0
16 to 24 years	17,282	32.6	41.3
25 to 54 years	78,682	41.0	43.3
55 years and over	11,692	36.7	42.3
Men, 16 years and over	57,362	42.1	44.5
16 to 24 years	8,989	34.7	42.3
25 to 54 years	42,124	44.1	44.9
55 years and over	6,250	39.6	43.7
Women, 16 years and over	50,294	35.8	40.8
16 to 24 years	8,293	30.4	40.0
25 to 54 years	36,558	37.4	41.0
55 years and over	5,442	33.3	40.3
Race and Hispanic origin			
White, 16 years and over	90,997	39.3	43.2
Men	49,114	42.4	44.8
Women	41,883	35.6	40.9
Black, 16 years and over	12,162	38.3	41.2
Men	5,826	40.0	42.3
Women	6,336	36.7	40.1
Hispanic origin, 16 years and over	9,645	38.5	41.5
Men	5,688	40.5	42.4
Women	3,956	35.6	39.9

Table 2 Average Weekly Hours of Work for Men and Women, 1976 and 1993

	Average hours		Age-adjusted hours
	1976	1993	1993
Men, 16+ years	41.0	42.0	41.2
Women, 16+ years	34.0	36.0	35.0

wage and salary workers who reported that they were at work exactly 40 hours per week declined, while the share working 49 hours or more rose. (A more detailed discussion of the shift among workers into the long-hours worked category is presented later in the section "Long Workweeks.") The proportions working fewer than 40 hours and 41–48 hours remained fairly stable.

Age and Sex

Since the changing age distribution affects workweek trends, it is useful to look at more homogeneous groups of workers over time. Between 1976 and 1993, the average workweek for 25- to 54-year-old men and women were both up on net. The increase was much greater for women, whose average workweek rose by nearly two-and-a-half hours (Figure 2). During that 17-year period, however, the workweek fluctuated substantially with the business cycle. Men's hours were curtailed more severely in conjunction with the downturn of the early 1990s, and, even by 1993, they had not yet regained their prerecession peak. Adult women, in contrast, experienced only a small dip in their average workweek, and that series quickly returned to its upward trend.

The slight increase in average hours worked between 1976 and 1993 reflects the greater share of both men and women who worked 49 hours or more per week (Table 3). For men, there was a corresponding decline in the share who worked exactly 40 hours per week, while among women the shift into the longer workweek occurred from the part-time category (1–34 hours) and from the 35- to 39-hour group.

Younger Workers

In contrast to workers aged 25 to 54, the average workweek for those who are younger edged down, on net, between 1976 and 1993.

Figure 1 Distribution of Hours at Work of Nonagricultural Wage and Salary Workers, Annual Averages, Selected

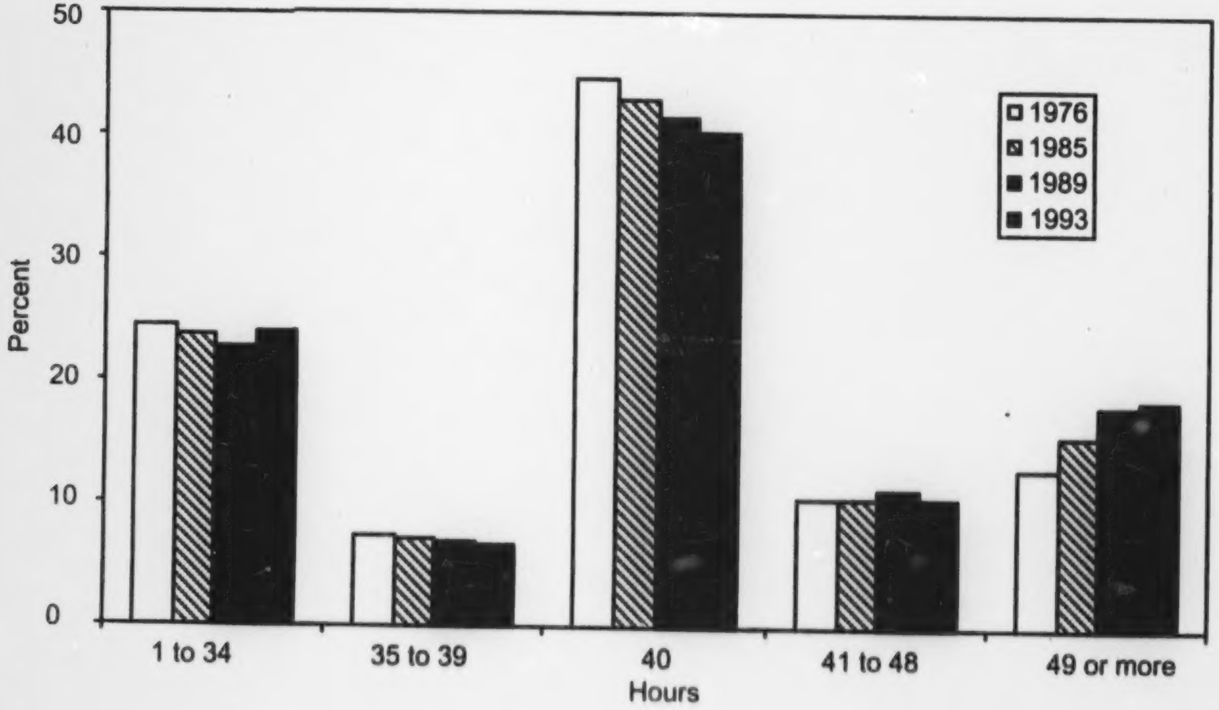
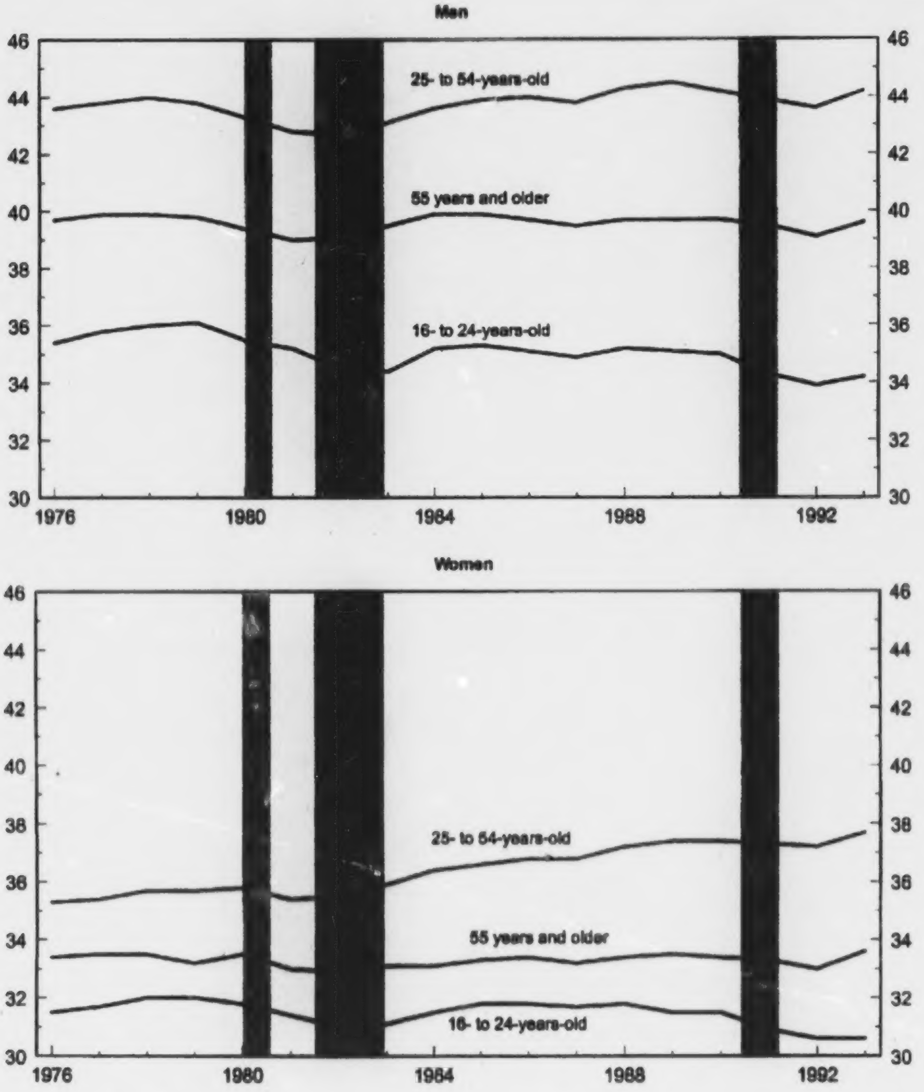


Figure 2 Average Hours Worked for Wage and Salary Workers in Nonagricultural Industries, by Sex and Age, Annual Averages, 1976–1993



NOTE: Shaded areas represent recessions.

Table 3 Percent Distribution of Nonagricultural Wage and Salary Workers, by Sex, Age, and Hours of Work, Annual Averages, Selected Years, 1976–1993[a]

Characteristic	1976	1985	1989	1993
Men				
16 to 24 years				
1–34 hours	34.1	35.9	36.7	40.2
35–39 hours	5.1	5.4	5.6	6.2
40 hours	38.5	36.3	35.6	33.0
41–48 hours	11.3	9.9	9.3	8.2
49 hours or more	11.1	12.6	12.8	12.4
25 to 54 years				
1–34 hours	10.4	9.8	9.1	10.7
35–39 hours	4.3	4.2	4.0	4.1
40 hours	48.9	46.6	43.7	42.7
41–48 hours	14.2	13.8	14.2	13.3
49 hours or more	22.2	25.7	29.0	29.2
55 years and over				
1–34 hours	18.3	19.1	21.4	23.0
35–39 hours	4.7	5.0	4.9	4.6
40 hours	50.7	46.6	43.5	41.9
41–48 hours	11.5	11.2	10.6	9.9
49 hours or more	14.7	18.1	19.7	20.6
Women				
16 to 24 years				
1–34 hours	43.3	44.5	46.1	50.5
35–39 hours	9.8	9.1	8.4	8.1
40 hours	37.8	34.1	32.8	29.4
41–48 hours	5.9	6.9	6.7	5.9
49 hours or more	3.2	5.3	6.0	6.0
25 to 54 years				
1–34 hours	31.4	28.2	26.1	26.5
35–39 hours	11.6	10.5	9.7	9.4

Characteristic	1976	1985	1989	1993
40 hours	43.8	43.5	43.3	42.4
41–48 hours	7.5	8.9	9.9	9.8
49 hours or more	5.7	8.9	11.0	12.0
55 years and over				
1–34 hours	38.4	39.4	39.5	40.4
35–39 hours	11.8	11.5	11.4	9.9
40 hours	38.5	37.5	35.3	35.2
41–48 hours	6.5	6.0	6.7	6.5
49 hours or more	4.9	5.6	7.1	7.9

[a] Detail may not sum to 100.0 due to rounding.

In 1976, 16- to 24-year-olds worked an average of 33.6 hours a week compared to 32.5 hours in 1993. While average hours at work were higher for young men than for young women (34.2 and 30.8, respectively, in 1993), the cyclical and long-term trends were nearly identical.

The overall decline in hours worked among youth partly reflected changes in their school enrollment status. As shown in Table 4, between 1976 and 1993, the proportion of all 16- to 24-year-olds who were attending school—either high school or college—increased from 45 percent to 50 percent. The rise in school enrollment occurred among both high school and college-aged youth.[9]

In addition to rising enrollment rates among the college-age population, more college students were working in 1993 than in 1976—53 versus 45 percent. This rise in employment occurred entirely among full-time college students, who averaged about 20 hours a week. Thus, the shift toward shorter workweeks among the young largely reflects their increased tendencies to be students. However, even among nonstudents, average hours edged down slightly.[10]

Table 4 School Enrollment of 16- to 24-Year-Olds in the United States (%)

			College	
Year	Total	High School	Part time	Full time
1976	45.2	24.9	3.0	17.3
1993	50.0	23.7	4.4	21.9

Hours distribution data reinforce the contention that the decline noted in the average workweek among younger workers is due, in part, to an increase in school activity. The proportion of younger workers who work part time (1–34 hours per week) has increased since the mid 1970s, while the share of those working 40 hours per week declined.

Older Workers

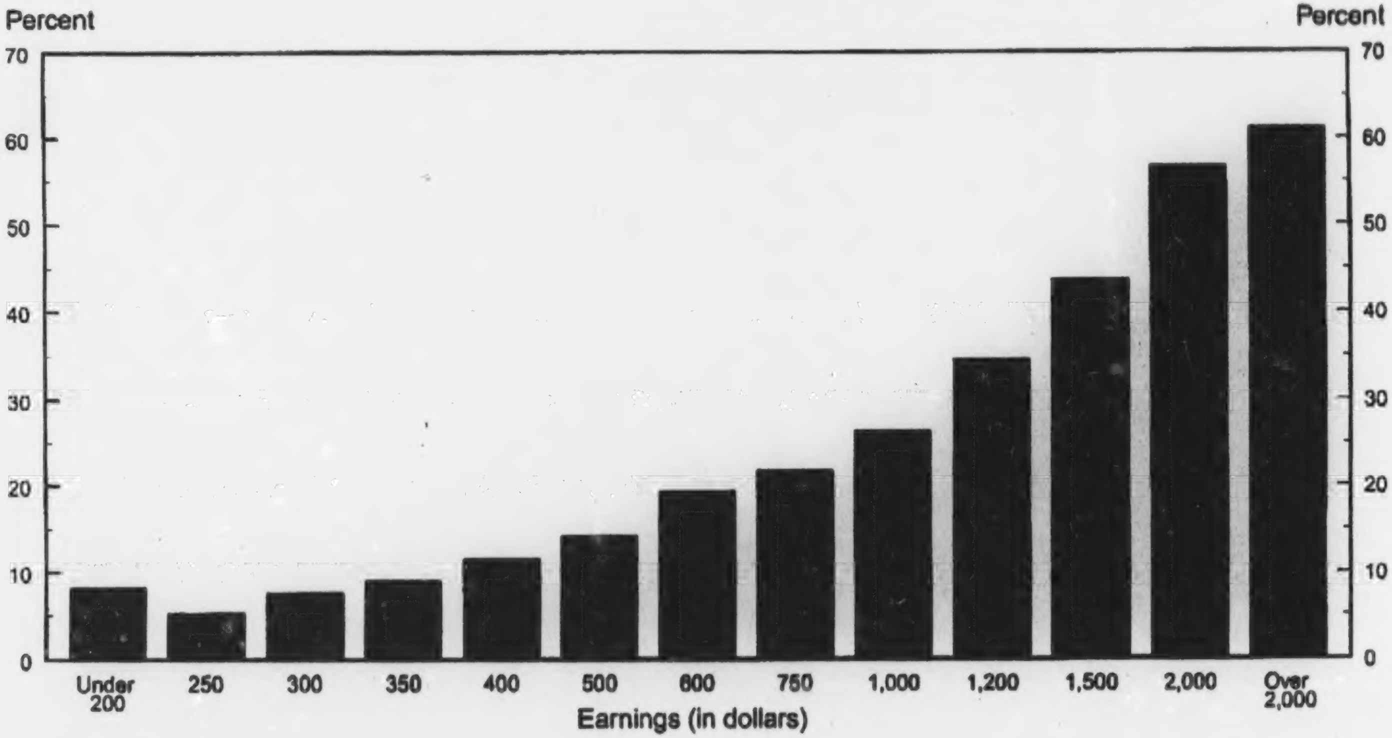
As with workers aged 25 to 54, men aged 55 and over work an average of about six hours more per week than their female counterparts. The average workweek for both men and women 55 years and older changed little between 1976 and 1993, and their averages seem to have been less affected by the business cycle than were those for other age groups. For older men in particular, the unchanged average workweek, on net, reflects increases in employment at both ends of the hours distribution (Table 3). Apparently, a growing share of those still in their "career jobs" are working very long workweeks, as was the case for workers aged 25 to 54. At the other end of the hours distribution, work activity among retirees (those receiving pensions) is on the rise, and these workers tend to work part time.[11] In fact, between 1984 and 1993, the proportion of pension recipients who worked rose from 31 to 39 percent.

LONG WORKWEEKS

Who Is Working Longer Workweeks?

It is a simple arithmetic truth that persons who work longer workweeks earn more, on average, at equivalent hourly pay, than those who work shorter workweeks. For example, persons working 48 hours per week at $10 per hour would earn $80 more, before taxes, than those working 40 hours per week at the same hourly rate. In addition, survey data from the CPS clearly show those with the highest earnings are quite likely to work very long hours[12] (Figure 3). What is not obvious from mathematical computations and survey data is which comes first: do the high earnings associated with longer workweeks simply reflect the greater hours worked, or is there a more basic difference between

Figure 3 **Proportion of Full-Time Men in Each Earnings Category Who Work 49 Hours or More Per Week, 1995 Annual Averages**



NOTE: Intervals reflect the upper bounds of the earnings categories.

jobs that demand (or encourage) long workweeks and those that do not?

Figure 4 shows the share of workers in different occupations who work 49 hours or more per week in 1985 and 1993. Professionals and managers are among those most likely to work very long workweeks. This may reflect the considerable responsibilities associated with many of these types of jobs, but also the fact that employers often are not required by law to pay them overtime premiums, as they must do for most hourly paid workers. Workers in these occupations also are among the highest paid: professionals earned $682 per week and managers $675 in 1993, compared with the median for all occupations of $463.[13]

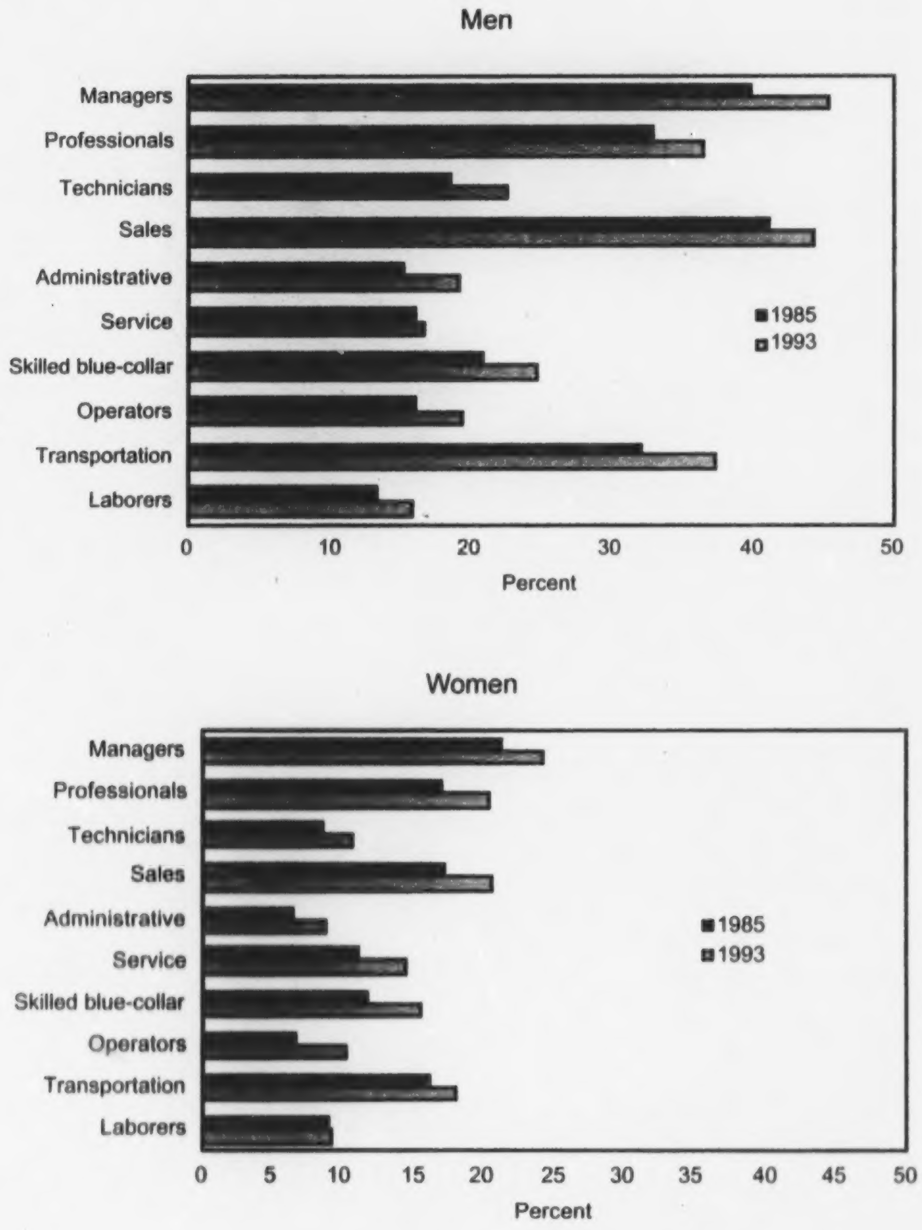
In contrast, sales and transportation workers also have long workweeks, but they are not, on average, highly paid. In these cases, a large percent of workers may work 49 hours or more a week due to the direct effect on earnings—that is, the more they work, the more they earn. For example, commissioned sales workers clearly have an incentive to work long workweeks. Indeed, full-time workers employed by motor vehicle dealerships averaged nearly 47 hours per week in 1995. In contrast, workers in department stores, where commissions are a less common form of pay, worked less than 41 hours.[14] Likewise, in the transportation industry, both trucking and taxicab services have among the longest workweeks of any industry, averaging more than 47 hours each.

To better understand the link between hours, occupations, and earnings, data for more specific occupations need to be examined. For example, even within the occupational groups where the overall share of workers employed 49 hours or more is small, there may be some types of jobs in which such schedules are common. Such an analysis, however, is outside the scope of this overview. The discussion presented here suggests that there are several factors that distinguish occupations with long workweeks, and that these jobs may be intrinsically different from other types of jobs.

1985–1993 Occupational Shift

Does the increasing share of workers who report that they are at work for more than 48 hours reflect a shift in employment toward high-hour occupations? For both men and women, the share in every major

Figure 4 Share of Workers on Full-Time Schedules Working 49 Hours or More Per Week, by Occupation, 1985 and 1993 Annual Averages



occupational group that worked such a schedule increased between 1985 and 1993[15] (Figure 4). As stated above, the prevalence of long workweeks varies considerably by occupation. Such schedules are more highly concentrated in the managerial, professional, sales, and transportation occupations, and the rate of increase during the 1985–1993 period was not consistent among all occupations. The following tabulation shows the distribution of growth in long-workweek employment across the occupational mix effect, the within-occupation shift effect, and the employment growth effect.[16]

As Table 5 shows, the number of persons working long work schedules increased considerably (5.1 million) over the eight-year period. Nearly half of this gain (2.4 million for both sexes combined) can be attributed to the overall expansion in employment over the period—the employment growth effect. The shift into occupations in which long workweeks are the most prevalent—such as managers, professionals, sales, and transportation—accounted for about 400,000, or 8.1 percent, of the gain for men and women combined. This occupational mix effect, however, was much larger for women than men, 12.7 versus 5.1 percent. The rest of the increase is due to the rise in the share of long workweeks in every occupation for both men and women, shown as the within-occupation shift effect.

Table 5 Growth in Employment of Persons Working 49+ Hours per Week, by Reason

	Total (000)	Men (000)	Women (000)
Number at work 49 hours or more			
1985	16,787	13,006	3.781
1993	21,909	16.093	5,816
1985–1993 change	+5,122	+3,087	+2,033
Occupational mix effect	+416	+158	+258
Within-occupational shift effect	+2,341	+1,259	+1,082
Employment growth effect	+2,365	+1,670	+695

NONAGRICULTURAL SELF-EMPLOYED AND AGRICULTURAL WORKERS

Although a growing share of nonagricultural wage and salary workers have long workweeks, most still have a workweek that is fairly close to 40 hours. In contrast, the majority of the self-employed have either very short or very long workweeks (Table 6). The proportion of the self-employed who work at least 49 hours per week declined between 1976 and 1993—although it is still nearly double that for nonagricultural wage and salary workers—while the share who worked part time (1–34 hours per week) rose. In contrast to the trend for men, who comprise the majority of the self-employed, the proportion of women who work a longer workweek increased since the mid 1970s, and the share working 1–34 hours per week declined. As with the self-employed, agricultural workers are heavily concentrated at both ends of the hours distribution and their share of workers in the 49+ hours

Table 6 Percent Distribution of Persons at Work, by Class of Worker and Hours of Work, 1976 and 1993 Annual Averages[a]

	Hours of work				
Class of worker	1–34	35–39	40	41–48	49 or more
1976					
Nonagricultural workers[b]					
Wage and salary	24.5	7.3	44.6	10.6	13.0
Self-employed	27.4	4.4	22.8	9.0	36.4
Agricultural workers	30.7	4.8	14.4	8.2	42.0
1993					
Nonagricultural workers[b]					
Wage and salary	24.0	6.7	40.3	10.6	18.5
Self-employed	31.0	4.9	23.3	7.0	33.9
Agricultural workers	29.3	4.8	22.1	7.6	36.2

[a] Detail may not sum to 100.0 due to rounding.

[b] Excludes unpaid family workers.

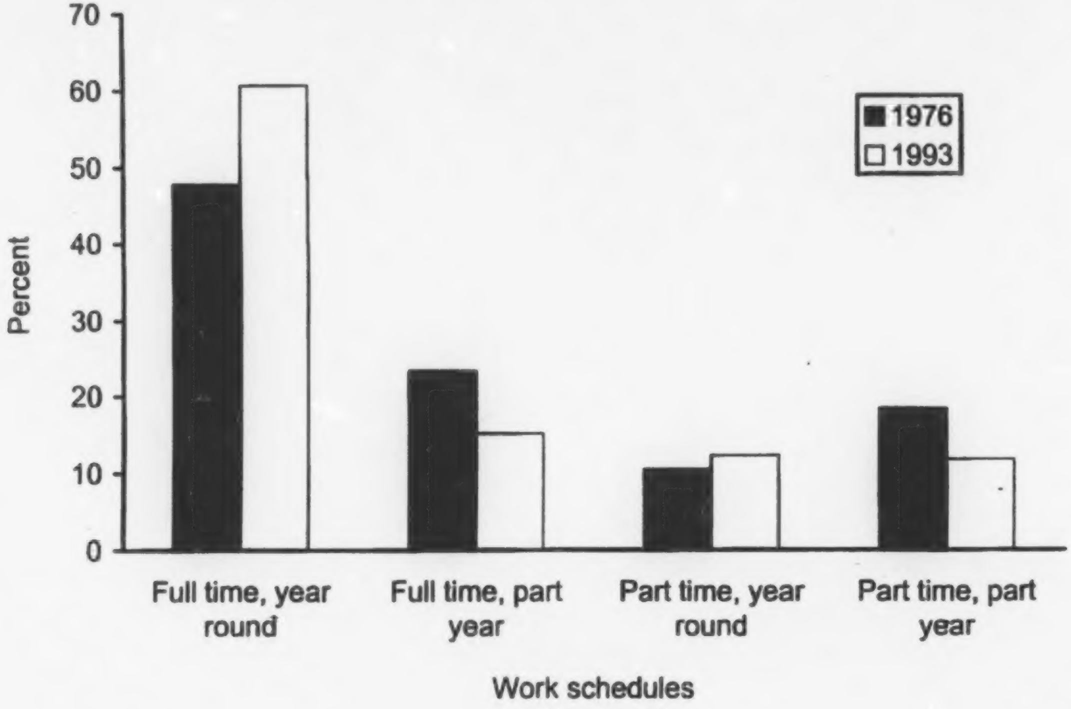
group declined substantially between 1976 and 1993, as the share working exactly 40 hours rose.

ANNUAL WORK HOURS

We have seen that the change in the average length of the workweek has been quite small since the mid 1970s, although a shift toward a growing share of workers putting in very long workweeks has been noteworthy. But rephrasing the question from, "What has been the trend in the length of the workweek?" to the broader, "What has been the trend in hours at work over an entire year?" brings in additional variables that may identify more dramatic shifts. Indeed, data on annual work hours, rather than the average workweek, most often are used in intercountry comparisons of work hours. This allows for the differences in vacation time allowed and used between, say, Germany, Japan, and the United States to be factored into the work-hours discussion.

Two factors other than the length of the typical workweek can affect the total amount of time people spend working: the extent to which people work at all during any particular year and the number of weeks that people work during the year. In the previous calculation of average weekly hours, workers are only included when they worked; they were "out of scope" when they did not work at all; that is, they are in neither the numerator nor the denominator of an average weekly hours calculation. Yet we know that changes have taken place in the amount of time during the year that workers are spending on the job. Bureau of Labor Statistics analysts recently reported that work activity is becoming less and less seasonal (more year round), and that finding is consistent across industries and demographic groups (see Rydzewski, Deming, and Rones 1993). Data collected each March in the CPS also show that U.S. workers, particularly women, have increasingly been working year round, as shown in Figure 5. Indeed, more dramatic than any shift toward either full- or part-time work is the trend toward year-round employment. The following shows the effect of changes in the share of the population working and the extent of their work activity during the year on work hours.

Figure 5 Work Schedules of Women Aged 25 to 54, 1976 and 1993 Annual Averages



The average number of hours the average worker is at work during the year is calculated using the following formula:

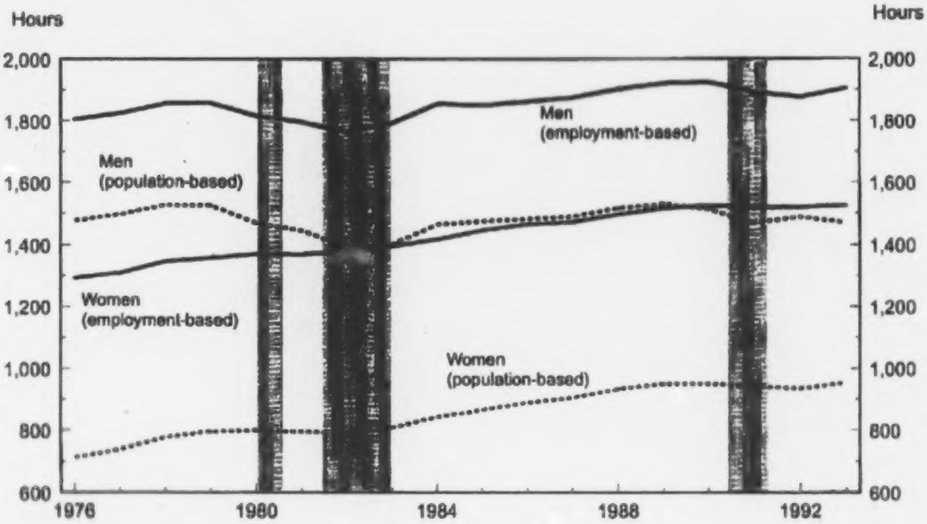
$$\begin{array}{c}\text{Average}\\ \text{annual hours}\\ \text{at work}\end{array} = \begin{array}{c}\text{Number at}\\ \text{work in an}\\ \text{average week}\end{array} \times \begin{array}{c}\text{Average}\\ \text{weekly hours}\\ \text{at work}\end{array} \times \text{52 weeks} \div \begin{array}{c}\text{Number at}\\ \text{work during}\\ \text{the year}\end{array}$$

The aggregate number of hours worked during a week is the product of the number of persons at work in an average week (an annual average) and their average hours at work. That number is then multiplied by 52 weeks to obtain an estimate of the aggregate number of hours worked during the year. The divisor—the number at work at any time during the year—was obtained from the "work experience" questions asked each March in the CPS supplement.[17] In those questions, CPS respondents are asked to recall their work activity during the previous calendar year, including the number of weeks in which they worked and their usual hours. Thus, aggregate hours worked during the year 1993, for example, obtained from the basic monthly CPS, are divided by the number of persons who worked at all in 1993 (that number is obtained from the March 1994 survey). This produces an excellent measure of average hours worked for each worker during the year and a long time series for comparisons. Results for men and women are shown in Figure 6.

The annual hours estimate rose steadily for women until the late 1980s and has grown more slowly since then. The lack of sensitivity to the business cycle is somewhat surprising given the fact that women, like men, are subject to cyclical swings in unemployment, which is a major determinant of the number of weeks worked during the year. The hours series for men is higher than that for women both because men work longer average workweeks and they are more likely to work year round. Men's annual hours have risen much less than women's and appear to be more sensitive to the business cycle.

As shown in Table 7, working women worked an average of nearly 20 percent more in 1993 than in 1976, adding 233 hours to their average workweeks. But, as shown with the weekly hours data, the age distribution of the working population has changed substantially over this period; a much smaller share are now in the older and younger ages,

Figure 6 Average Annual Hours at Work for Men and Women, 1976–1993



NOTE: Shaded areas indicate recessions.

where both workweek length and weeks of work tend to be relatively low. Adjusting for this age shift reduces modestly the 1976–1993 change. Men's hours, after age-adjustment, were up 3 percent over the period; women's, up 15 percent.[18]

These calculations still leave one important trend identified earlier unaccounted for: the change in the likelihood of an individual to work at all during the year. Men have become less likely to work, largely due to earlier retirements, expansion of the Social Security disability program, increased school enrollments, and an increase in wives' employment. Women, alternatively, have become dramatically more

Table 7 Average Annual Work Hours for Men and Women, 1976 and 1993

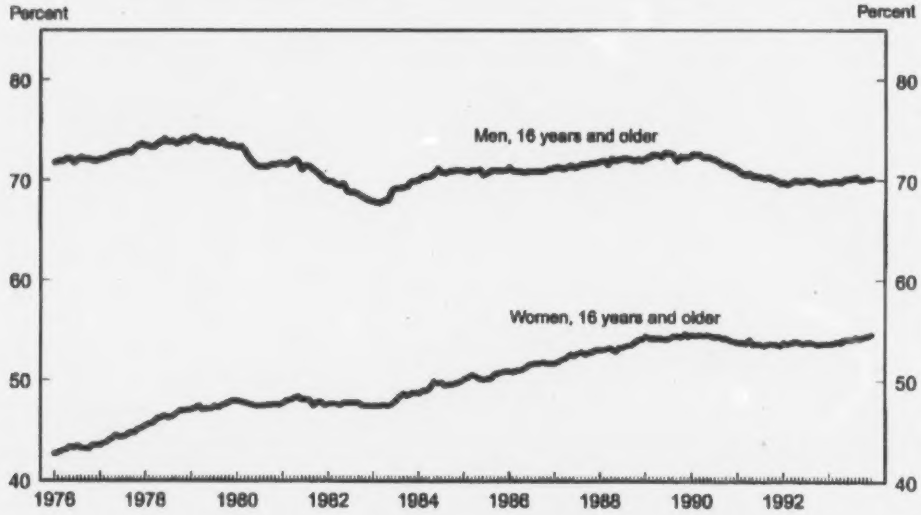
Year	Men	Women
1976	1,805	1,293
1993	1,905	1,526
1976–1993 change	+100	+233
Age-adjusted change	+62	+193

likely to work (Figure 7). Hence, using the population as the denominator in an annual hours calculation, rather than those who worked, should considerably affect the change in hours between 1976 and 1993.

The population-based estimate duplicates the numerator in the equation above but uses the civilian noninstitutional population as the denominator, not those who worked. As shown in Figure 6, the series for men did not change at all. In fact, the line is flatter than the employment-based series since men, on average, have become somewhat less likely to work at all over time. The population-based series for women is at a much lower level than is the employment-based series. The increases, though, have been quite large, particularly on a percentage basis. Allocated across the population of women aged 16 and older, each individual worked one-third longer in 1993 than in 1976.

Looking at the more homogeneous (in terms of work schedules) group of 25- to 54-year-olds has two advantages: it avoids the need to age-adjust the data, and it eliminates the younger and older workers from the calculation, the two groups with particularly low annual hours. For that group, between 1976 and 1993, average hours per woman per year rose 45 percent, from 888 to 1,290. The average for men was virtually unchanged at just over 1,900 hours (see note 18).

Figure 7 Employment-Population Ratios for Men and Women, Annual Averages, 1976–1993



SUMMARY

This chapter attempted to track the course of working hours in the United States using the Current Population Survey, a large, representative national sample of households from which comparable data can be obtained for a long period of time. The survey estimates suggest that the average length of the workweek for most groups has changed little since the mid 1970s, although the distribution of work hours has changed. The most noteworthy difference between the 1970s and the 1990s is in the increase in the share of persons who are working very long hours—those who are exceeding the old "standard" workweek of 40 hours by more than a full eight-hour day. This increase is pervasive among occupations, and the long workweek itself seems to be associated with high earnings and certain types of occupations.

More dramatic has been the increase in the work year, a measure more commonly used to compare different countries (Americans tend to work more during the year than most Europeans but less than the Japanese, for example). Women's increasing likelihood of working at all and, when they do work, of working year round, has had a noteworthy effect on the number of hours that women work during the course of the year. In order to analyze trends in either weekly or annual work hours over an extended period of time, it is important to allow for changes in the overall measures that occurred solely because of changes in the age distribution of the population. Alternatively, the analyst can examine trends for specific population groups separately.

Notes

1. See Ilg (1996) for a discussion of the industries and occupations that experienced job growth in recent years.
2. This trend analysis ends in 1993 due to the introduction of a redesigned Current Population Survey (CPS) in January 1994. The new CPS asked different questions to obtain average hours data from the pre-1994 survey, rendering the data not strictly comparable. See the appendix for a discussion of changes in the CPS and its effect on work hours. Data for 1995 are presented, however, in the overall description of between-group differences in work hours.
3. An additional limitation of the Current Employment Statistics survey, the Bureau of Labor Statistics' survey most commonly used for average workweek data, is that the universe is restricted to private nonsupervisory workers on nonagricul-

tural payrolls. The excluded groups—agricultural workers, the self-employed, and many supervisory and professional workers—tend to have very different levels of work hours than do those who are covered.

4. The CPS is a monthly survey of 50,000 (at present) households conducted by the Bureau of the Census for the Bureau of Labor Statistics. Another source of data on work time comes from time diaries. This approach is discussed in Robinson and Bostrom (1994).
5. The restricted group is presented because those excluded—nonagricultural self-employed and agricultural workers—have very different workweeks. Those differences are discussed later in the chapter. In addition, the workweek decisions are conceptually very different for the self-employed than they are for "employees," who must match their own preferences with those of employers.
6. Marital-status data are for all workers in nonagricultural industries, not just wage and salary.
7. In 1995, full-time workers aged 25 to 54 worked an averaged of 44.1 hours a week, about three hours longer than the average for all workers that age. The long-term trend in the workweek for full-time workers is similar to that for all workers; that is, fluctuating with the business.
8. To "age-adjust" the length of the workweek, first the age distributions of men and women at work in 1976 were applied to the 1993 employment total for each gender to generate a new 1993 distribution. Aggregate hours then were computed by multiplying the new employment figures for each age by the average hours worked in 1993. The aggregate hours for the age groups were then summed individually by sex to get total aggregate hours for men and women. These totals were then divided by male and female employment in 1993 to obtain an age-adjusted workweek that uses the age distribution of 1976 and the age-specific hours worked in 1993.
9. The share of 16- to 24-year-olds who are enrolled in high school appears to have fallen, according to the tabulation. That is so only because this population group has shifted substantially toward the college ages over that period. In 1976, 51 percent of 16- to 24-year-olds were teenagers; by 1993, that share was only 43 percent. In fact, the enrollment rate for 16- to 19-year-olds in high school was 48 percent in 1976 compared to 55 percent in 1993.
10. Hours data for nonstudents were available only for 20- to 24-year olds. In 1979, their average workweek was 40.4 hours compared to 39.7 hours in 1993.
11. See Herz (1995) for a discussion of several possible reasons for the increased work activity of pension recipients.
12. The data shown are for men but the relationship applies to women as well.
13. Earnings data presented are for full-time (35 hours or more a week), wage and salary workers. Earnings data are not available for self-employed sales workers or for those earning commission.
14. These data are for industries, not occupation; data on hours at work by detailed occupation are not produced regularly by the Bureau of Labor Statistics.

15. These dates were selected because the occupational classification system used prior to the early 1980s was quite different from the one put into place in the CPS in 1983. Data beyond 1993 were affected by the redesign of the CPS introduced in January 1994. Each year selected is more than two years after the end of the prior recession, so estimates of change should not be influenced by the business cycle. These data do include the self-employed.
16. The *employment growth effect* is a measure of the change that would have occurred simply as a result of the overall growth in employment. Thus, it gives the 49+ group its "fair share" of the overall 1985–1993 growth. The *occupational mix effect* is derived by estimating the number of persons who would have worked 49+ hours in 1993 if the occupational mix had been the same as it was in 1985. The *within-occupation shift effect* reflects the extent to which the change in 49+ employment over the period are due to the changes in the share in each occupation who work 49+ hours, as shown in Table 4. It applies the share in each occupation who worked such schedules in 1985 to the actual occupational employment distribution in 1993.
17. Such an estimate cannot be derived from the basic monthly CPS.
18. The basic CPS data include a break in the population (and employment) series between 1990 and 1991. Data from 1990 forward have been adjusted to 1990 census estimates, adjusted for the undercount. March work-experience data, however, have not been so revised. Thus, a slight inconsistency exists between the numerator and denominator in the average annual hours calculation when pre- and post-1990 data are used. The effect on the data is minimal, particularly when long-term comparisons such as the 17 years used here are made. See Robert J. McIntire, "Revisions in Household Survey Data Effective February 1996," *Employment and Earnings*, March 1996, pp. 8–14, for a discussion of the revisions to the population series.

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Appendix: Changes in CPS Questionnaire Concerning Hours Worked

Current Population Survey (CPS) data for January 1994 and forward are not strictly comparable with data for earlier years because of the introduction of a major redesign of the questionnaire and collection methodology. The principal reasons for the redesign were to obtain more accurate information on the labor market in general, and to expand the use of computer technology in the data collection process. Among the questionnaire changes were alterations to the questions on the number of hours actually worked during the reference week. The questions were modified to help respondents recall the exact number of hours they worked on their main jobs in the prior week. This appendix describes the differences in the questions asked to obtain hours-at-work data in the pre- and post-1994 surveys. In general, the changes emphasized the importance of precision in recalling the prior week's work activity, but they do not alter the concept of hours at work.

In an effort to obtain more precise hours-at-work data, beginning in 1994, respondents to the new CPS are first told that the following questions focus on the exact number of hours they worked in the prior week. They are then asked if they lost or took off any hours from their jobs for any reason in the prior week. If yes, they are queried about the number of hours. Respondents are also asked if they worked extra hours at their jobs that they do not usually work and, if so, how many. It is not until these prompts are completed that respondents are asked how many hours they actually worked at their jobs, and, in addition, for multiple jobholders, how many hours they actually worked at their other jobs.

Prior to 1994, the questions pertaining to actual hours were slightly different, as was the ordering of those questions (see questions below). Data on actual hours were obtained by first asking the number of hours worked at all jobs last week. Then questions were asked about taking time off and working extra hours. The onus was placed on the interviewer to correct the original answer of hours worked, if necessary, based on responses to these questions. Also, nothing in the interview communicated the importance of precision to the respondent. In the pre-1994 survey, hours data were collected for all jobs combined

Comparing pre- and post-1994 data suggests that the implicit recall strategy associated with the new questionnaire does provide more accurate data on actual hours (see Appendix Table 1). For instance, the proportion of persons who reported working exactly 40 hours per week—a common, almost reflex, response—declined substantially between 1993 and 1994. In fact, this decrease was greater than the cumulative effect of the long-term downward trend between 1973 and 1993. In addition, during the 1973–1993 period, the share of survey respondents reporting that they worked between 35 and 39 hours or 41 and 48 hours had decreased. In 1994, with the revised questions, this trend was reversed, indicating that respondents now are giving different, and apparently more precise, answers to the questions on hours actually worked.

The following questions were used to obtain data on actual hours worked in the new and old CPS:

New CPS	Old CPS
Lead-in: Now I have some questions about the exact number of hours you worked last week.	How many hours did you work last week at all jobs?
Last week, did you lose or take off any hours from (work/your main job), for any reason such as illness, slack work, vacation, or holiday?	Did you lose any time or take any time off last week for any reason such as illness, holiday, or slack work?
(If yes) How many hours did you take off?	(If yes) How many hours did you take off?
Last week, did you work any overtime or extra hours (at your main job) that you do not usually work?	Did you work any overtime or at more than one job last week?
(If yes) How many additional hours did you work?	(If yes) How many extra hours?
So, for last week, how many hours did you actually work at your (main) job?	Interviewers are instructed to correct original answer if lost time was not already deducted or if extra hours were not included.
(For multiple jobholders) Last week, how many hours did you actually work at your other job(s)?	

Table A1 Percent Distribution of Persons at Work, by Sex and Hours of Work, 1993 and 1994 Annual Averages[a]

Characteristic	1993	1994	Difference
Men (hours)			
1–4	0.4	0.7	0.3
5–14	2.6	2.4	–0.2
15–29	8.1	8.4	0.3
30–34	5.7	6.3	0.6
35–39	4.5	5.3	0.8
40	41.1	37.1	–3.9
41–48	12.1	14.3	2.2
49 or more	25.5	25.5	–0.1
Women (hours)			
1–4	0.8	1.1	0.4
5–14	5.1	5.4	0.3
15–29	16.5	17.3	0.9
30–34	9.8	10.2	0.4
35–39	9.2	10.2	1.0
40	39.4	35.1	–4.3
41–48	8.8	10.3	1.6
49 or more	10.5	10.3	–0.3

[a] Detail may not sum to 100.0 due to rounding.

3
Working Hard

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There is considerable variation in the hours worked and in the reported devotion to work of persons among advanced Organisation for Economic Co-operation and Development (OECD) countries. Americans, the Japanese, Australians, and New Zealanders put in lots of hours on the job. Relatively many Americans and Canadians report that they want to work more hours than they do. By contrast, western Europeans enjoy long vacations and considerable leisure while employed, and in many European Union countries work sharing is encouraged as a method for dealing with unemployment. During the 1980s and 1990s the gap between time worked by employed Americans and Canadians and time worked by their western European comrades increased noticeably.

There is also considerable divergence in earnings inequality among advanced OECD countries, with inequality higher in the United States, Canada, and the United Kingdom than in most OECD-Europe countries. During the 1980s and 1990s, moreover, inequality also grew much more rapidly in the United States, the United Kingdom, and to a lesser extent Canada than in continental western European countries (Freeman and Katz 1994).

To what extent, if at all, are these two patterns related? Does high earnings inequality induce workers to work longer hours and work harder? Has increased inequality contributed to the rising gap in time worked between workers in the United States and Canada and those in Europe? How does the greater work time of Americans affect compar-

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isons of economic performance between the United States and other OECD countries?

This chapter presents the basic facts about "working hard" in the United States and Canada relative to other advanced OECD countries. It sketches out the hypothesis that inequality in outcomes increases work activity and offers some preliminary evidence from the United States regarding this hypothesis. The empirical evidence reveals a positive relationship between the hours worked within detailed occupation, industry, and region cells, and the inequality in hourly wages in those cells is consistent with the hours-inequality hypothesis.

NORTH AMERICANS AS WORKAHOLICS

Hours Worked and Preferences

Table 1 presents estimates of annual hours worked and changes in annual hours worked in major advanced OECD countries. Column 1 records annual hours worked per employed person as reported by the OECD. The sample of employees includes part-time as well as full-time workers. Annual hours are higher in the United States than in the major European countries, although workers put in many hours in several other countries as well, most notably Japan, Australia, New Zealand, and Finland. Hours worked by employed Canadians are 3 percent lower than hours worked by employed Americans, but they are still above the hours worked in most advanced OECD-European countries. Annual hours per employed person does not, however, capture the full difference in working time among countries because there are also sizable differences in the ratio of employees to the adult population, due in part to labor force participation decisions and in part to differences in rates of unemployment across countries. In 1994, for example, the employment/population ratio for 16–64 year olds was 73.2 in the United States, 64.2 in Canada, and 58.2 in OECD-Europe (OECD 1995). Column 2 of Table 1 records employee/population ratios for the various countries for which we have annual hours data. Multiplying the annual hours per employed person by the employment/population ratios gives the annual hours worked per person of working

Table 1 Differences in Annual Hours Worked among Advanced OECD Countries, 1994

Country	Annual hours per employee, 1994	Employment population ratio (ages 15–64)	Annual hours per adult (ages 15–64)	Estimated change in annual hours per employee, 1970–1999[a]
United States	1,780	73.2	1,303	–121.5
Canada	1,719[b]	63.8	1,097	–148.5
United Kingdom	1,717[b]	66.5	1,142	–120.8
Norway	1,415	72.7	1,029	–346.0
Sweden	1,544	70.3	1,085	–65.9
Germany	1,578	62.6	988	–389.0
Finland	1,780	60.1	1,070	–204.0
France	1,631	59.0	962	–320.4
Netherlands	1,395	63.7	889	–361.3
Australia	1,882	67.0	1,261	—[c]
New Zealand	1,843	68.2	1,257	—
Japan	1,965[d]	74.2	1,458	–236.0[e]

SOURCE: Column 1, OECD (1995, Table C); column 2, OECD (1995, Table A); column 4, calculated from OECD (1996, Table 3).

[a] Based on estimates of annual average change in hours from trough to trough over three time periods, as given in OECD Employment Outlook, 1996. Exact time periods by country as follows: 1) United States, 1970–71; 2) Canada, 1970–92; 3) United Kingdom, 1971–93; 4) Norway, 1970–92; 5) Sweden, 1972–93; 6) Germany, 1971–94; 7) Finland, 1971–93; 8) France, 1971–93; 9) Netherlands, 1972–93.

[b] 1993 data.

[c] A dash implies that data for these countries were not provided in the OECD table.

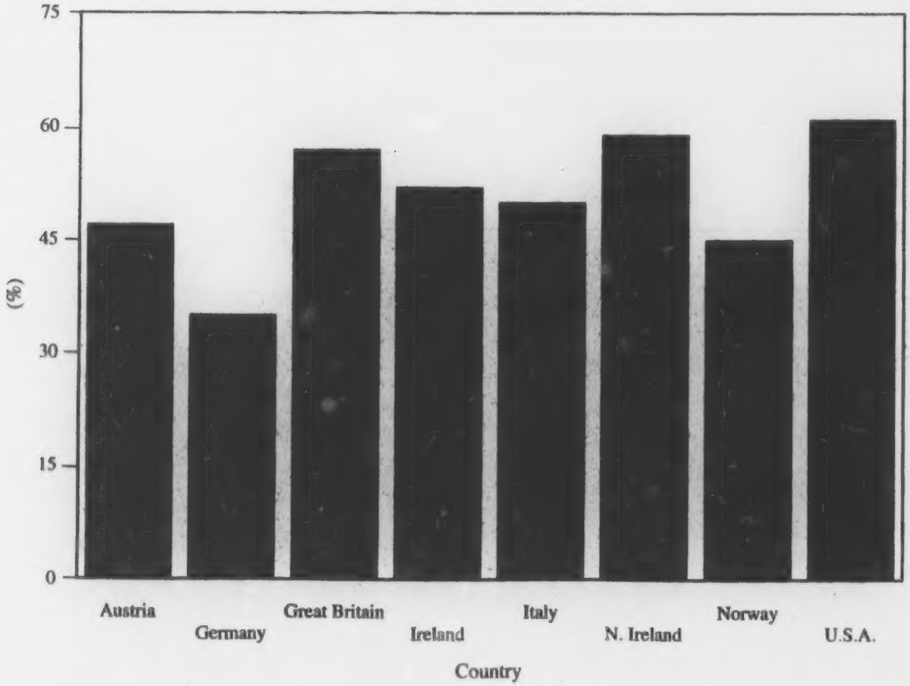
[d] 1992 data.

[e] Based on change in actual hours per employee (1973–93), as reported in OECD Labor Force Statistics.

age in column 3. The differences between the United States and western Europe in hours worked per adult are on the order of 30 percent, whereas those between Canada and western Europe are relatively modest. To the extent that hours worked per adult are a measure of "working hard," Americans work harder than Canadians and western Europeans.

There are two additional pieces of evidence that support the claim that Americans are more devoted to work than are western Europeans. In its 1989 World of Work module, the International Social Science Programme (ISSP) survey[1] of workers in different countries contained the following question: Which of the following statements best describes your feeling about your job? 1) I work only as hard as I have to, 2) I work hard but not so much that it interferes with the rest of my life, 3) I make a point of doing the best work I can, even if it interferes with the rest of my life. Figure 1 compares the proportion of workers who gave

Figure 1 Percentage of Workers Who Work Hard "Even if It Interferes with the Rest of Their Life"



SOURCE: Bell and Freeman (1995, Table 5.7).

the third response—working hard even if it interfered with their lives—among the advanced OECD countries covered by the 1989 survey. The figure shows that U.S. workers were the most likely to work hard at the expense of the quality of their lives, followed by persons in other English-speaking countries. Germans, Norwegians, and Austrians, on the other hand, were the least likely to sacrifice for their jobs.

The second piece of evidence comes from surveys that ask individuals to choose between working more or fewer hours than they currently do. These questions are a bit tricky, because by specifying the hypothetical differently, one can readily induce different but valid responses. We focus on questions that ask people about the desire to work fewer or more hours at the same rate of pay, as opposed to questions that ask about the desire to work more hours at an overtime rate or about preferences between increases in pay for the same hours of work versus the same pay for reductions in hours worked.[2]

For the United States, data on preferences come from the May 1985 Current Population Survey (CPS) Supplement. The specific question analyzed is: If you had a choice would you prefer to work: 1) the same number of hours and earn the same money, 2) fewer hours at the same rate of pay and earn less money, or 3) more hours at the same rate of pay and earn more money? For Canada, the June 1985 Canadian Labour Force Survey asked a more detailed and complicated question that also specified that the employees would be paid the same rate, while at the same time indicating that all other conditions of work remained the same (see Kahn and Lang 1988). For European countries, the March 1991 *European Economy* reports results from a European Economic Community (1991) survey that asked the question this way: Assuming that your present hourly rate remained unchanged, would you like to work less, as long, or longer? For Japan, the 1992 Employment Status Survey (*Shugyo Kozo Kihon Chosa*) asked a question comparable to the May 1985 CPS question.

Table 2 summarizes the results from these diverse surveys. It shows a striking difference in preferences for more or less work between Americans and Canadians and western Europeans, and between Americans and Canadians and the Japanese as well. While in all countries the majority of people are satisfied with their current hours at work, the proportion wanting to work more hours than they currently do is higher for Canadians and Americans than for Europeans

Table 2 Preferences for More or Fewer Hours of Work

Country	Prefer more hours and more earnings	Prefer same hours and same earnings	Prefer fewer hours and less earnings	Differences between columns 1 and 3
Canada (1985)	35	50	15	20
United States (1985)	27	65	8	19
Japan (1992)	3	68	30	–27
Germany (1989)	4	55	38	–34
United Kingdom (1989)	4	65	30	–26
Europe	9	51	37	–28

SOURCE: Canada—tabulated from data in Kahn and Lang (1988).

United States–May 1985 Current Population Survey, as reported in Bell and Freeman (1995).

Germany– European Economic Community (1991, Table 2).

United Kingdom–British Social Attitudes Survey.

Europe–European Economic Community (1991, Table 22). Data include Belgium, Denmark, Italy, Netherlands, Portugal, and the United Kingdom. In the E.U. study, U.K. figures are 12 percent for more hours/earnings, 50 percent for the same, and 12 percent for less, giving a difference of –21.

or Japanese, and the proportion who want to work fewer hours is lower for Americans and Canadians than for Europeans and Japanese. The differences in preferences among countries are well-summarized by the final column of Table 2, which shows the differences among countries between the proportions of individuals wanting more and less work. The fact that North Americans work more hours than Europeans and at the same time have a greater preference for additional hours worked than Europeans is particularly noteworthy. By contrast, the Japanese, who also work many hours, want to work fewer hours than they currently do and, in preferences if not actual hours worked, more closely resemble the Europeans.[3]

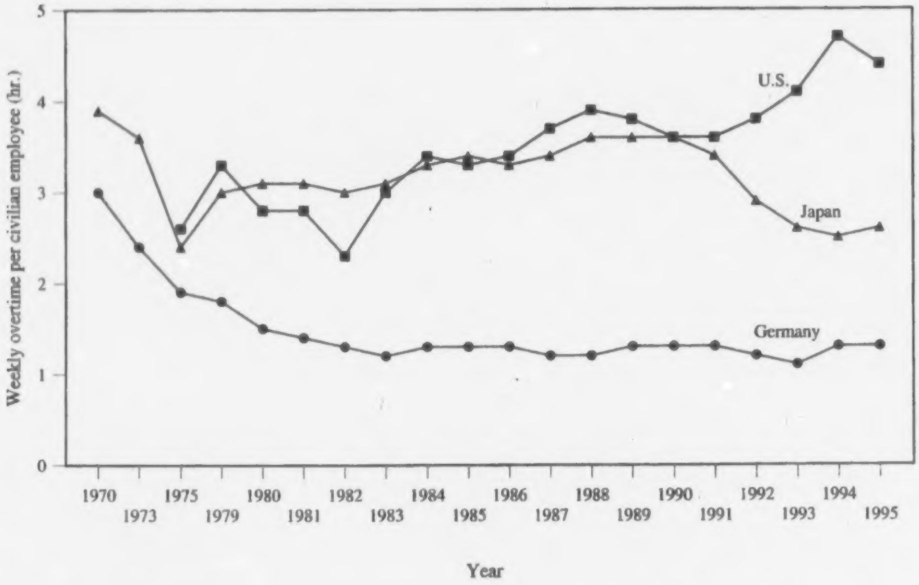
In sum, on the basis of all three statistics—hours worked, willingness to sacrifice for work, and desire to work more—Americans appear to be working harder than Europeans. As in many other statistics, Canada falls somewhere between the United States and Europe, showing high preferences for additional hours of work in surveys of preferences but not actually working all that much more than Europeans.

Changes in Hours over Time

Has the hours difference between North Americans and western Europeans always existed, or is this a relatively recent phenomenon? The evidence in column 4 of Table 1 shows that the greater work activity by North Americans developed in the 1970s–1980s. According to the OECD estimates, in 1970 (adding column 4 to column 1), North Americans worked fewer hours than Europeans. This finding is not unique to the OECD data; it is found in other statistics as well. Data gathered by Maddison (1995) for instance, show a similar pattern in hours worked per capita from 1950 to 1992. In 1950, Americans worked 24 percent fewer hours per capita than Germans, 18 percent fewer per capita than the French, 15 percent fewer than the British, and 8 percent fewer than Italians. By 1973 these differences narrowed greatly, as Europeans took much of their increased prosperity in leisure. By 1992, Americans worked 6 percent more hours than Germans, 22 percent more than the French, and approximately 12 percent more than the British or Italians. Between 1950 and 1992, hours worked per person in the United States was roughly constant, while hours worked per person in Europe fell by 17 percent (Italy) to 33 percent (France). Data from Japan provided annually by the Japan Productivity Center also show a drop in hours worked per employee of 4 percent from 1980 to 1991 compared to an increase in hours worked per employee in the United States of 11 percent. The Japan Institute of Labor (1994–1995) reports a fall in hours actually worked, including overtime, from 203 hours per month in 1960 to 159 per month in 1993—a 22 percent fall. While most hours series still show that the Japanese work more hours than Americans, the once-immense hours gap has diminished greatly. In Japan, the decline in hours is presumably linked in part to changes in national legislation intended to reduce hours to a 40-hour workweek by 1997 (OECD 1996), and for this reason it will likely continue.

Evidence on the amount of overtime hours worked—for which covered U.S. workers receive time-and-a-half overtime pay and for which workers in other countries often receive less premium—also shows a trend upward in U.S. overtime hours versus Germans and the Japanese (Figure 2). Whereas in 1994 overtime hours in the United States were at a post–World War II peak, overtime hours in Germany

Figure 2 Trends in Overtime Hours



SOURCE: OECD (1996, Table 3.7).

were much below those in the early 1970s. Overtime hours in Japan were also considerably lower than in the 1970s, with the most significant overtime hour declines in the last several years. The OECD data in Figure 2 show Americans with the highest amount of overtime of the three countries. While data from the Japanese Institute of Labor show that Japanese workers still put in more overtime than Americans, it also confirms that the difference in overtime hours has diminished greatly with the trend downward in overtime hours in Japan.

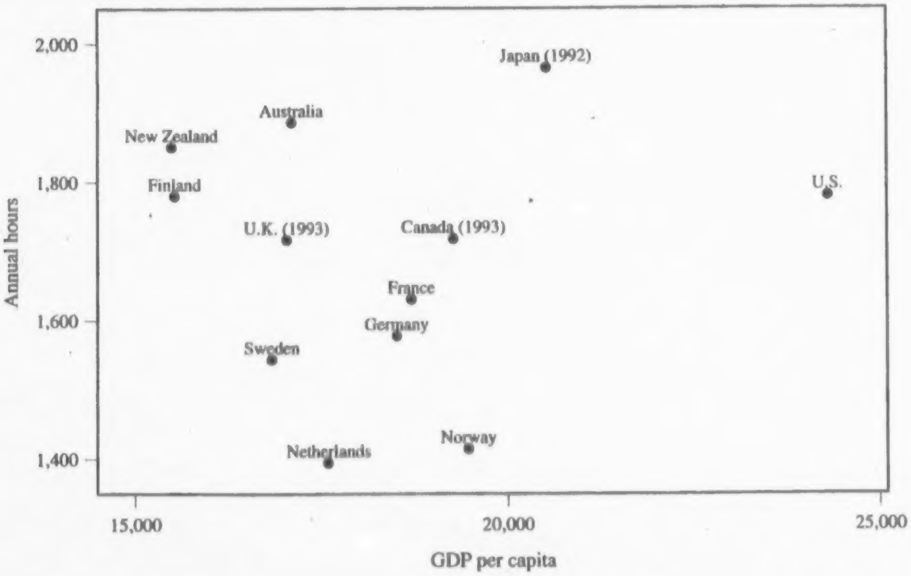
The fact that North Americans worked more hours than western Europeans in the 1990s and worked fewer hours than Europeans in the 1970s makes it difficult to explain cross-country differences in work time in terms of diverse culture or national psychology. The fact that the increased preferences of North Americans for greater work relative to Europeans seems to be a recent phenomenon (Bell and Freeman 1995) supports this as well. Instead of focusing on cultural differences, we direct our attention to differences and changes across countries in the economic incentives that induce workers to work many hours.

Hours Worked and GDP Per Capita

How does working hard relate to national income per capita? From a production function perspective, one might expect additional employment per adult (more properly, employment per capita) to be associated with higher GDP per capita—more input means more output. From a labor demand perspective, one might also expect a positive GDP per capita/labor input association: if higher GDP per capita reflects higher capital per capita, this would produce greater demand for labor and thus greater employment per capita. But a labor supply perspective suggests the opposite: falling time worked with higher income due to the income effect. Indeed, the labor supply–driven story is the usual one given for the long-term downward trend in hours worked, and would seem to fit the drop in hours in western Europe post–World War II.

Figure 3 rejects the notion that either production/demand or supply side forces dominate the relation between hours worked by employees

Figure 3 Annual Hours Worked per Employee vs. GDP Per Capita, by Country



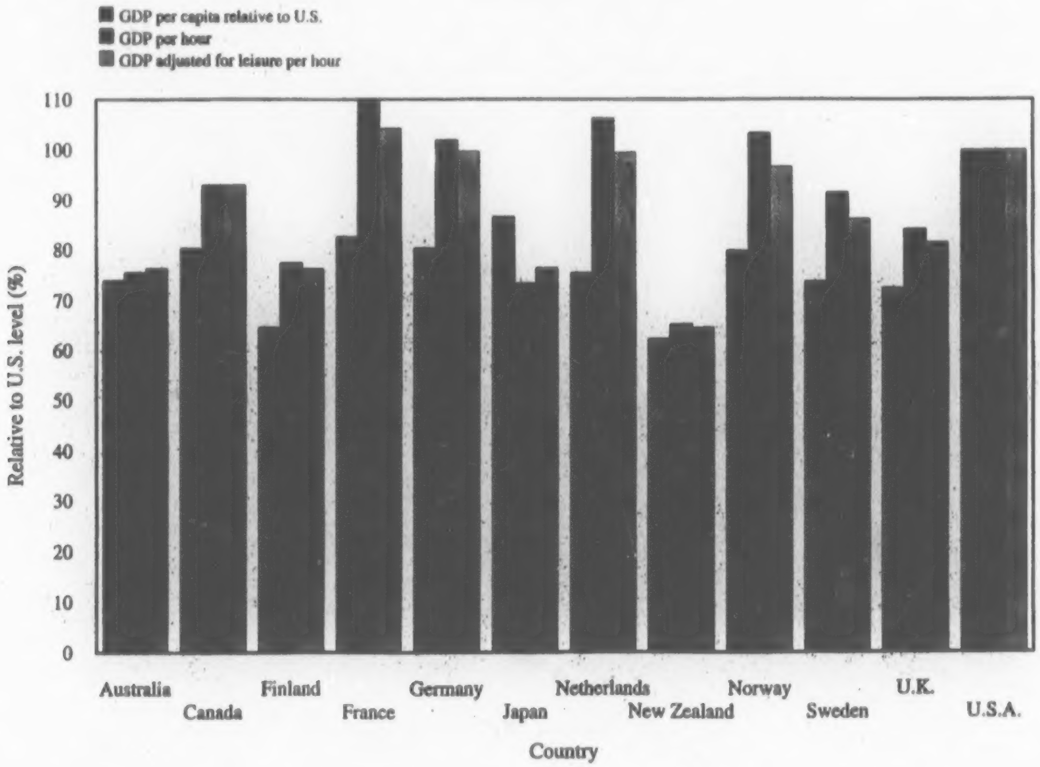
SOURCE: Annual hours, OECD (1995, Table C). GDP per capita from U.S. Bureau of the Census, Statistical Abstract 1995, Table 1374.

and income per capita (measured in purchasing power parity) among countries. There are high per capita income countries where employees put in lots of hours, such as the United States and Japan, and low per capita income countries where employees also work many hours, such as New Zealand. While Canadians and Norwegians have similar per capita incomes, they have very different annual hours worked, and although Germans have a relatively high income per capita, they work relatively few hours. Indeed the message from Figure 3, if any, is that an English heritage (save for Japan), not national income per capita, determines worktime across countries.

The lack of any clear relation between hours worked and GDP per capita notwithstanding, the wide variation in annual hours worked per adult among advanced OECD countries (Table 1, column 2) suggests that GDP per capita, unadjusted for differences in work time among countries, may be a seriously misleading indicator of national productivity and well-being in cross-country comparisons. On the productivity side, GDP per capita will understate the productivity of labor in countries where adults work fewer hours and overstate it in countries where adults work more hours. GDP per working hour arguably offers a better measure of productivity (Freeman 1995), although it is by no means perfect.[4] Measured by GDP per hour worked as opposed to GDP per capita, the sizable lead that the United States has in national productivity diminishes greatly. In GDP per capita, for instance, in 1993 the United States had a 31 percent advantage over Germany. In GDP per hours worked, by contrast, the United States and Germany had virtually identical productivity (Figure 4). But since Americans are working more hours than employees in these countries, we are potentially further down the marginal product of labor curve and thus are probably still more productive. Without measures of capital and other input, it is not possible to go much beyond the basic statement that the United States isn't as far ahead of others as initially appears to be the case.

On the welfare side, fewer hours worked by the employed implies greater leisure, which presumably adds to a worker's utility. Similarly, persons who choose not to participate in the workplace produce valuable goods and services at home and/or enjoy greater leisure. The nonwork hours of the unemployed, by contrast, is more difficult to assess: with good benefit programs and high reservation wages, one cannot

Figure 4 GDP Per Capita, GDP Per Hours, and GDP Adjusted for Leisure



SOURCE: Calculated as described in the text with hours per adult from Table 1; GDP per capita in purchasing power parity units from the U.S. Bureau of the Census (1995, Table 1374); adults per capita from OECD, Historical Statistics (1992, Table 2.1). Adjusted GDP per capita assumes valuation of Leisure = GDP/Hours Worked in U.S.

value unemployed hours at zero, but evidence that the unemployed are less "happy" than others (Clark and Oswald 1994) clearly implies that their time should be given a lower valuation than that of others. In any case, standard neoclassical analysis suggests that adults in countries with fewer hours worked will be better off relative to those in countries with more hours worked at the same level of GDP per capita.

In Figure 4, we pursue this logic by adjusting GDP per capita (measured in purchasing power parity units) for differences in hours worked per adult among countries. We take U.S. GDP per capita and hours worked per adult as the numeraire and estimate a leisure-augmented GDP per capita for other countries (*x*), based on their hours worked versus those in the United States using the following formula:

$$\begin{aligned}(\text{Aug. gdp/cap})_x = \text{gdp/cap}_x + (&\text{leisure hrs}_{x(\text{per adult})} \\ &- \text{leisure hrs}_{us(\text{per adult})}) \\ &\times (\text{adults per capita}_x)(\text{valuation of leisure})\end{aligned}$$

The difficult component of the equation is the value attached to the greater leisure of adults in other countries versus the United States. One possible valuation is to set the value of leisure at GDP per hour worked in a country, and in this case the equation simplifies nicely to:

$$(\text{Aug. gdp/cap})_x = \text{gdp/cap}_x\ [1 + (\text{workhrs}_{us} - \text{workhrs}_x)/\text{workhrs}_x]$$

Another alternative is to value leisure in country *x* at GDP per hours worked in the United States, giving us a more conservative estimate of augmented GDP per capital, which takes U.S. work hours as the base. This valuation simplifies to

$$(\text{Aug. gdp/cap})_x = \text{gdp/cap}_x\ [1 + (\text{workhrs}_{us} - \text{workhrs}_x)/\text{workhrs}_{us}]$$

In Figure 4 we report the results of the more conservative calculation. We take GDP/capita in purchasing power parity units from the U.S. Bureau of the Census (1995, Table 1374). We estimate the additional leisure that adults in country *A* have versus the United States by taking the difference between hours worked per 15- to 64-year-old in the United States and hours worked per 15- to 64-year-old in country *X* from Table 1 of this chapter. We obtained values for adults/capita from OECD Historical Statistics (1992), table 2.1.

Even with this conservative estimate of the value of leisure, our leisure augmented GDP per capita substantially compresses the position of the United States as the top country in the OECD league tables and considerably lowers the standing of Japan. More sophisticated analysis, valuing the nonwork hours of different people differently, would presumably produce somewhat different estimates but in the same direction, reducing the position of the North American hard-working countries. Because we work so hard, standard GDP per capita country comparisons indicate that we aren't as well off as our European compatriots.

Factors in Employee Work Time

What factors underlie the gap in hours worked or its complement, hours of leisure per employee, among advanced OECD countries?

The gap in hours worked and thus in hours of leisure per employee between the United States/Canada and western Europe may be due to the potential contribution of three factors, namely, differences in the proportion of workers who are part-time, differences in weeks of vacation (and holiday to a much lesser extent) time, and differences in hours worked per week by full-time workers.

Table 3 records the proportion of workers in various countries who are part-time. Even though European labor markets are less flexible along some dimensions than North American labor markets, the proportion of jobs held by part-timers is higher in many European countries than in the United States, although it is lower in both Germany and France than in the United States. Indeed, part-time working stands out as the sole major form of nonstandard working that has shown a substantial increase since 1970 in the majority of European OECD countries (OECD 1996).

Given reasonable estimates of the difference in hours worked between part-time and full-time workers, however, it is difficult to explain much of the hours gap among employees between the United States and other countries in terms of part-time work, even for the European countries with very high part-time rates. Consider, for example, the case of the Netherlands, where 35 percent of workers are part-

Table 3 The Role of Part-Time Work in Hours Worked in Advanced OECD Countries, 1994

Country	Share of part-time workers (%)	Overall yearly change in work (hrs.)	Change attributable to: Change in share of part-time workers (%)	Change attributable to: Change in work of full-timers (hrs.)	Change attributable to: Change in work of part-timers (hrs.)
United States	18.9	—[a]	—	—	—
Canada	17.0	—	—	—	—
Belgium	12.8	–7.5	–4.9	–2.5	0.2
Denmark	23.3	–6.6	1.4	–7.1	–0.9
Italy	6.2	–3.7	–0.9	–3.0	0.4
United Kingdom	23.8	–1.5	–0.5	3.8	–0.5
Norway	26.5	—	—	—	—
Sweden	24.9	—	—	—	—
Germany	15.1	–10.9	–3.9	6–.1	–0.9
Finland	23.3	—	—	—	—
France	14.9	–4.1	–4.4	0.4	0.7
Netherlands	35.0	–6.6	–11.3	0.0	3.2
Australia	24.4	—	—	—	—
New Zealand	21.6	—	—	—	—
Japan	21.4	—	—	—	—

SOURCE: Column 1, OECD (1995, Table E; 1993 for Canada, the United Kingdom, and Denmark; 1992 for Japan). Columns 2–5, OECD (1996, Table 3.2).

[a] A dash implies that data for these countries were not provided in the OECD table.

time. The hours gap among employees in Table 1 between the United States and the Netherlands is 485 hours. Assume that the hours worked by part-timers are 60 percent of those worked by full-timers. Let *dP* be the difference between the proportion of Dutch and American workers who are part-time; according to Table 3, this is 16 percentage points (35–19 percent). Then, if the Dutch had the same proportion of part-time workers as Americans, their average hours worked would increase by $0.4/[1 - 0.4\,(0.35)]\ dP$, or by 7.4 percent.[5] This would bring Dutch hours to 1,498, closing the hours gap by 103 hours, or 21 percent. For the other countries, the effect of increasing the part-time proportions to U.S. levels would be markedly smaller, while for Germany and France, the calculation works in the opposite direction: the low levels of part-time work imply that the adjustment would increase the difference in hours worked. The vast bulk of the difference in annual hours worked between North Americans and Europeans is evidently attributable to differences in the hours of full-time workers.

While part-time work cannot explain much of the U.S.–Europe hours gap, the increase in part-time work does help explain the 1980s–1990s fall in hours worked among European countries. Column 2 of Table 3 records the overall average yearly change in annual hours worked in European economies from 1983 to 1993 (with slight variation among countries due to differences in the data). Columns 3–5 decompose that change into the part due to changes in the share of part-timers in employment and in the hours worked of full-time and part-time workers. There is wide variation in the decomposition. In France all of the 1983–1993 drop in hours is due to an increased share of part-timers. In Germany, by contrast, the bulk of the decline in annual hours is due to falling hours of full-time workers. In the Netherlands, the increase in part-timers "overexplains" the fall in hours; the compensating factor is an increase in the hours of part-timers. In the United Kingdom, the hours worked by full-time employees works in the opposite direction to the change in part-timers. Additional data for other European countries also show considerable variation in the importance of part-time work to changes in annual hours.

If differences in the pattern of part-time work among countries do not explain the bulk of the U.S.–European work hours differences, what does? Table 4 reverts back to this issue by considering the contribution of differences in weekly hours and vacation/holiday time in explaining the hours of full-time workers. The data in Table 4 relate to full-time employees in manufacturing because that is the only sector for which we have readily available internationally comparable data. However, scattered information for workers in other sectors (for instance, from the Union Bank of Switzerland study of prices and earnings around the globe) tells a similar story. Indeed, because many countries legislate vacation time or determine it through national collective bargaining, differences among sectors within a country tend to be modest.

Column 1 of Table 4 gives the annual hours worked of full-time manufacturing workers in the various countries. Column 2 of the table records the amount of vacation and holiday time in each country measured in five-day weeks. Column 3 gives standard hours per working week and is measured exclusive of overtime hours. Note that there are substantial differences in vacation and holiday time, due almost entirely to vacations—in the United States the typical worker has a 2.4 week vacation compared to 5.1 weeks for the typical European. Hours worked per week differ much less, with Germany and Norway having the lowest scheduled hours.

Column 4 calculates the difference in annual hours worked between each of the countries and the United States. The differences are large: the 261-hour difference in annual hours of full-time workers in manufacturing between the United States and Germany is 6.5 full

Table 4 The Contribution of Vacation/Holidays and Weekly Hours to Differences in Annual Hours Worked per Full-Time Employees in Manufacturing

				Differences U.S.A. (hr.)		
Country	Annual work (hr.)	Vacation/ holiday (5-day weeks)	Work per week (hr.)	Overall	Due to vacation/ holidays	Due to hours
United States	1,904	4.6	40.0			
France	1,763	7.0	39.0	141	94	48
Germany	1,643	8.5	37.6	261	147	114
Italy	1,764	8.1	40.0	140	140	0
Netherlands	1,709	8.3	38.9	195	144	52
Norway	1,718	6.4	37.5	186	68	119
Sweden	1,784	7.6	40.0	120	120	0
United Kingdom	1,769	6.6	38.8	135	78	57
Europe average (unweighted)	1,736	7.5	38.8	168	113	57

SOURCE: Tabulated from Bell and Freeman (1995, Table 5.2), using data from the Federation of German Employer's Association.

weeks of work, and the 120-hour difference between the United States and Sweden is 3 full weeks of work. Column 5 gives our estimate of the contribution of vacation and holiday time to the difference in annual hours between the United States and other countries. Column 6 gives our estimate of the contribution of hours worked per week to the difference in annual hours between the United States and other countries. In both of these calculations we take the United States as the base and calculate the annual hours worked in other countries as if they had U.S. vacation and holiday time, or as if they had U.S. weekly hours.

The numbers in columns 5 and 6 show that much of the observed difference in annual hours worked between the United States and other countries is attributable to the low vacation and holiday time in the United States. With the sole exception of Norway, the annual hours difference due to differences in vacation and holiday time are larger than the differences due to hours worked. In the case of Italy and Sweden, where scheduled hours are the same as in the United States, all of the difference is due to vacation and holiday time. In Germany and Norway both vacation and holiday time and hours worked per week contribute substantially to the annual hours gap with the United States. For the other countries, the differences in vacation and holiday time dominate the observed difference between U.S. annual hours and the country's annual hours.

The final line in Table 4 presents a crude summary of the factors underlying country work hour differences. It gives unweighted averages of annual hours, vacation and holiday time, and hours worked for the European countries covered in the table. It also shows the contribution of vacation and holiday time and hours worked to the difference between the average annual hours and hours in the United States. Approximately two-thirds of the gap between annual hours of full-time workers is attributable to differences in vacations and holidays and one-third to differences in hours per week.

WHY WORK SO HARD?

Standard labor supply analyses link individual work hour decisions to wages and nonlabor income. In standard analysis, changes in market wages have an ambiguous effect on work time or effort because these changes have both an income and substitution effect typically illustrated in textbooks with an indifference curve diagram. Only in the case of the pure substitution effect can responses to changes in wages be signed, holding income/utility fixed, wage changes should induce individuals to substitute hours in the same direction as the change in wages. Nonlabor income has an unambiguous effect on work hours, with increases in income reducing time worked if leisure is a normal good. Changes in nonmarket productivity or wages will also have an unambiguous effect on time worked in this model, since higher nonmarket opportunities increase total income and induce substitution of work time to nonmarket time. Note that in standard labor supply presentations, inequality of earnings opportunities does not enter the supply decision in any obvious way. Instead, the standard model focuses on the effect of a change in individual wages without considering changes in the distribution of wage opportunities in the market.

In marked contrast, analyses of labor supply concerned with designing contracts to motivate workers place great stress on the shape of the opportunities frontier facing workers and thus on the distribution of opportunities. Piece rate or incentive pay schemes link rewards to effort measured in terms of output. Tournament pay systems link rewards to relative effort. In linking hours worked to the dispersion of opportunities in the relevant market, our analysis of differences in hours worked across countries or among persons in different markets within a country builds on the insights from these types of models.

Consider, for example, two workers, each of whom faces a differently shaped earnings opportunities set due to differences in the distribution of pay in the labor market, differences in job security provisions, or differences in unemployment insurance or other safety net provisions. Hans works in Germany, where pay differences among firms or within a firm among workers are relatively modest, where there is considerable job security, and where unemployment benefits in any event are high and relatively long-lived. Hank works in the United

States, where there are large pay differences among firms or within a firm among workers, where employment at-will produces a high degree of job insecurity, and where unemployment benefits are more modest and relatively short-lived. Who is more likely to work more hours and put in more effort on his current job? If Hans doesn't work that hard he doesn't lose all that much, and if he works hard he doesn't gain all that much either. But if Hank doesn't work hard he can lose his job and suffer painful unemployment or a sizable fall in pay at a new job. On the other hand, if he works hard, Hank can rise in the highly unequal pay distribution and make much more money.

Expressed differently, if the percentile position of a worker in the earnings distribution in his market (either through the firm that employs him, promotions within that firm, or pay within a job grade in that firm) depends on his hours worked/work effort, greater inequality in pay will induce greater work effort. For U.S.–Europe contrasts, our hypothesis can be decomposed into three steps:

1) For an incremental hour of work/effort, employees improve themselves in the relevant earnings distribution commensurately in terms of percentile position in the United States and Europe (if U.S. workers' earnings rise more,[6] this simply augments our story).

2) Any given change in the distribution of earnings translates into a larger difference in earnings in distributions with greater dispersion of pay, and therefore American absolute earnings are more dependent on percentile position than European absolute earnings.

3) Individuals respond to differences in the return to hours/effort with greater hours/effort.

Expressed somewhat differently, the hours-inequality argument is that a mean-preserving spread of wages raises effort/hours. The correct incentive variable in a labor supply equation is not the current wage (as in many labor supply analyses), but the incremental change in the lifetime-expected stream of income due to an increment in effort/hours today—the derivative in lifetime income streams with respect to an additional hour/effort at work. Because we believe that this derivative is positively affected by pay inequality, we expect higher inequality to be associated with greater hours/effort. To the extent that the level of the wage an individual receives affects the percentile position and therefore the expected return to hours/effort, wage levels matter as well.

Empirical Evidence

Is pay inequality, in fact, related to hours worked?

Bell and Freeman (1995) showed a positive rank correlation between the variance of ln (earnings) and mean weekly hours among full-time workers across nine countries in the 1989 ISSP survey, which is suggestive of just such a relation. Specifically, using data on earnings and hours worked from nine countries including the United States, Germany, the United Kingdom, Netherlands, Austria, Italy, Ireland, Northern Ireland, and Norway, the correlation analysis performed by Bell and Freeman (1995) showed a strong association between the hours-worked ranking of a country and the variation in ln earnings ranking in that country, but no significant association between the hours-worked ranking and the mean-earnings ranking, as might follow from standard labor supply analysis.

Using data from the May 1985, 1989, and 1991 CPSs, we build on evidence of within-cell variation among occupations, industries, and regions in hours worked and test the role of wage variation in explaining these hours patterns.

Specifically, we grouped workers into categories of noncompeting markets by detailed industry, detailed occupation, detailed industry-occupation, detailed industry-region, and detailed occupation-region as defined within the CPS. The rationale for this decomposition is to arrive at labor market cells that reasonably contain the distribution of wages relating to an individual worker's future opportunities. Exploiting the fact of significant hours differences across cells, we attempt to explore the role of differences in the derivative of lifetime opportunities with respect to hours/effort in explaining the hours patterns. Absent such measures, we estimate the incentive for workers to put in more hours/effort by the dispersion of pay in the job market in which they work.

Throughout the bulk of our analysis, we concentrate on full-time, private nonagricultural workers (working 35+ hours).[7] Using the two-digit categorizations of industry and occupation in the CPS and limiting the data to private nonagricultural workers gave us 42 potential two-digit industries and 41 potential two-digit occupation cells to use for our calculations. The hours figures used in this analysis are "usual hours worked per week," as reported in the CPS. In calculating the hourly earnings of nonhourly workers, we divided usual weekly earnings by usual hours worked. The hourly earnings of hourly workers are self-reported in the CPS.

For each detailed industry and/or occupation-region cell, we calculated four statistics:

1) the mean hours worked in the relevant cell,

2) the mean ln (hourly earnings),

3) the standard deviation in ln (hourly earnings), and

4) the 90/10 percentile ln earnings spread of full-time workers.

Appendix Table A1 shows the calculated statistics of mean hours, pay, and inequality in pay by industry for each year. Appendix Table A2 shows the resultant estimates by occupation for each year. We note sizable variation in mean hours worked across both industry and occupation in these tables.

Table 5 summarizes the basic relationship in these data in terms of the correlation coefficients between hours worked and the level and dispersion of pay across the relevant cells. In each of the three years analyzed, we obtain a positive and in most cases significant correlation between hours worked and the dispersion in hourly earnings, measured by either the standard deviation or the 90/10 percentile spread in ln (hourly wages). The strong positive relationship between hours worked and wage variability is more robust than the cross-section relation between hours worked and the level of pay—work hours are higher in higher paid occupations but lower in higher paid industries in two of the three years.

How robust is the empirical relation between inequality of pay and time worked? As shown in Table 5, changes in the measure of inequality of pay do not noticeably affect the relationship nor do changes in the cell categories.[8] However, in order to provide a further check on the basic relationship, we regressed hours worked by an individual in

Table 5 The Correlation between Hours Worked and the Level of Variance of Wages, by Detailed Industry, Occupation, and Region Cells as Indicated[a]

Cell category	Mean ln (hourly earnings)	Std. ln (hourly earnings)	90–10 spread
May 1985 CPS Data			
Detailed industry *n*=41	–0.220	0.200	0.111
Detailed occupation *n*=42	0.380**[b]	0.399**	0.411**
Detailed industry—occupation *n*=721	0.105**	0.061	0.072
Detailed industry—region *n*=160	–0.249**	0.096	0.166**
Detailed occupation—region *n*=161	0.179**	0.328**	0.346**
May 1989 CPS Data			
Detailed industry *n*=42	0.240	0.118	0.035
Detailed occupation *n*=43	0.425**	0.457**	0.285
Detailed industry—occupation *n*=829	0.214**	0.082**	0.000
Detailed industry—region *n*=164	0.145**	0.240**	0.194**
Detailed occupation—region *n*=166	0.293**	0.224**	0.261**
May 1991 CPS Data			
Detailed industry *n*=42	–0.120	0.225	0.239
Detailed occupation *n*=42	0.249	0.591**	0.486**
Detailed industry—occupation *n*=866	0.158**	0.070**	0.070**
Detailed industry—region *n*=167	–0.199**	0.080	0.143
Detailed occupation—region *n*=164	0.178**	0.541**	0.266**

[a] For private nonagricultural workers, 35+ hours.

[b] ** Indicates statistical significance at greater than 0.05% level.

the 1991 May CPS sample on a set of measures of personal characteristics, family income, and the wage, together with our measures of market inequality: the standard deviation of ln (hourly earnings) in an occupation-industry cell; the standard deviation of ln (hourly earnings) in an occupation cell; and the standard deviation of ln (hourly earnings) in an industry cell. Table 6 records the results of these calculations. The results are clear: all of the measures of inequality are estimated to have a positive effect on hours worked. Since (as is common in cross-section calculations like these) family income obtains a positive coefficient while the hourly wage has a negative coefficient in the hours regression, we would not interpret the equation as a labor supply relation, but rather as a check on the robustness of the inequality-time worked correlation that we argue is a more appropriate measure of the incentive to work hard than standard wage measures.

While we regard the results from Table 5 and 6 as supportive of the hours-inequality hypothesis, we note that this result is sensitive to one change in specification, namely, the inclusion of part-time workers in the sample. With part-timers included, the significant positive correlation between inequality of pay and hours worked disappears. Among industries, the correlation became negative in two of our three years whereas among occupations it remains positive but insignificant. One reason for this pattern is that there is considerable measurement error in the pay and possibly hours of part-timers that produces a large standard deviation in pay. Part-time work is associated with spurious inequality in pay, and, by definition, with fewer hours worked in a cell. This will bias the correlation of inequality to hours downward. Appendix Table A3 shows, however, that even with this bias, excluding part-timers who report working less than 10 hours per week, we obtain regression results with part-timers that are weaker than those in Table 6 but still support the basic hours-inequality hypothesis.

The hours-inequality hypothesis would, of course, be strengthened by evidence from other countries that workers respond to the incentives brought about by greater pay inequality by changing their actual work hours, as well as evidence that pay inequality affects desired work hours. While we lack detailed household data from other countries, the 1989 ISSP data allow us to evaluate worker preferences across countries. Ideally, we would want to analyze individual responses to two types of questions: how important individuals believed hours/effort to

Table 6 Hours Regressions, May 1991 CPS[a]

Dependent variable: ln (usual hours worked per week)

Independent variables										
ln (hourly wage)	–0.013	–0.020	–0.007	–0.020	–0.019	–0.016	–0.009	–0.014	–0.011	–0.007
	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
ln (family income)	0.011	0.082	0.010	0.008	0.008	—	0.009	0.009	0.009	—
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)		(0.002)	(0.002)	(0.002)	
Std. dev. of ln (hourly wage) (ind. occ. cell)[b]	0.159	0.323	0.103	0.020	—	0.019	—	—	—	—
	(0.013)	(0.014)	(0.014)	(0.015)		(0.015)				
Std. dev. If ln (hourly wage) (occ. cell)[c]	—	—	—	—	—	—	0.402	0.326	0.327	0.333
							(0.021)	(0.021)	(0.023)	(0.022)
Marriage dummy	—	0.011	0.017	0.012	0.011	0.012	—	0.015	0.015	0.015
		(0.003)	(0.003)	(0.003)	(0.003)	(0.003)		(0.003)	(0.003)	(0.003)
Female dummy	—	–0.051	–0.052	–0.049	–0.049	–0.048	—	–0.056	–0.051	–0.050
		(0.003)	(0.003)	(0.003)	(0.003)	(0.003)		(0.003)	(0.003)	(0.003)
Less than high school	—	–0.032	–0.014	–0.005	–0.007	–0.007	—	–0.013	–0.014	–0.015
		(0.008)	(0.008)	(0.008)	(0.007)	(0.004)		(0.007)	(0.007)	(0.007)
Some high school	—	–0.010	–0.011	–0.010	–0.010	–0.011	—	–0.012	–0.011	–0.013
		(0.007)	(0.007)	(0.007)	(0.007)	(0.006)		(0.007)	(0.007)	(0.006)
Some college	—	0.002	0.005	0.003	0.003	0.004	—	0.001	0.003	0.004
		(0.005)	(0.005)	(0.005)	(0.005)	(0.005)		(0.005)	(0.005)	(0.005)

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College graduate	—	0.026 (0.004)	0.034 (0.004)	0.028 (0.004)	0.027 (0.004)	0.027 (0.004)	—	0.026 (0.004)	0.029 (0.005)	0.029 (0.004)
College plus	—	0.053 (0.006)	0.071 (0.006)	0.056 (0.006)	0.055 (0.006)	0.053 (0.006)	—	0.054 (0.005)	0.061 (0.006)	0.059 (0.005)
Union dummy	—	–0.012 (0.004)	–0.023 (0.004)	–0.015 (0.004)	–0.015 (0.004)	–0.017 (0.004)	—	–0.016 (0.004)	–0.020 (0.004)	–0.022 (0.004)
Region dummy	No	Yes	Yes	Yes	Yes	Yes	No	Yes	Yes	Yes
Industry[d] dummy	No	No	Yes	Yes	Yes	Yes	No	No	Yes	Yes
Occupation[d] dummy	No	Yes	No	Yes	Yes	Yes	No	No	No	No
R^2	0.0287	0.139	0.110	0.150	0.151	0.147	0.050	0.103	0.125	0.122
N	8,615	8,615	8,615	8,615	8,820	9,926	8,820	8,820	8,820	9,445

[a] For private, nonagricultural workers, 35+ hours. Standard errors are in parentheses.

[b] Standard deviation in ln (hourly wage) in detailed industry/occupation cell.

[c] Standard deviation in ln (hourly wage) in detailed occupation cell.

[d] Detailed occupation and industry dummy variables.

be in determining pay/promotion/status within their work, and an individual's current hours and desired hours of work.

The 1989 ISSP data offer some valuable insight in this regard. Workers were asked to comment on "how important the quality of their work" was in determining their pay. An individual had the choice to respond in one of three ways: very important, somewhat important, or not very important. To the extent that the quality of an individual's work reflects work effort or hours, as it reasonably might, our theory would predict that individuals who respond that quality is extremely important would work more hours. As Figure 5A makes clear, we find precisely this relationship in the data for all countries. Figure 5B cross-tabulates responses to this question and to the question on how hard an individual works. Note that individuals who believe quality is extremely important in determining pay are more likely to "work hard even if it interferes with the rest of their lives" than are workers who believe quality to be relatively unimportant. Once again, with the exception of Austria, these results are uniformly true among countries in the sample. In sum, evidence from the ISSP offers additional support to the hours-inequality hypothesis with respect both to actual and desired hours of work, and suggests that workers in other countries respond similarly to workers in the United States in their hours preferences.

CONCLUSION

This chapter has documented the fact that Americans, and to a lesser extent Canadians, work hard, putting in more hours—and wanting to put in even more—than employees in many other advanced countries. It has shown that taking account of differences in the amount of time worked across countries alters the position of the United States in standard comparisons of GDP per capita. It has presented calculations of the relationship between inequality of pay and hours worked within U.S. occupation and industry job markets that suggest inequality of pay contributes to hours worked. In sum, in the United States we work hard because we face a good "carrot" for putting out time and effort, and because we also face a substantial "stick" if we do not.

Figure 5A Average Weekly Hours and Feelings about Importance of Work Quality for Pay

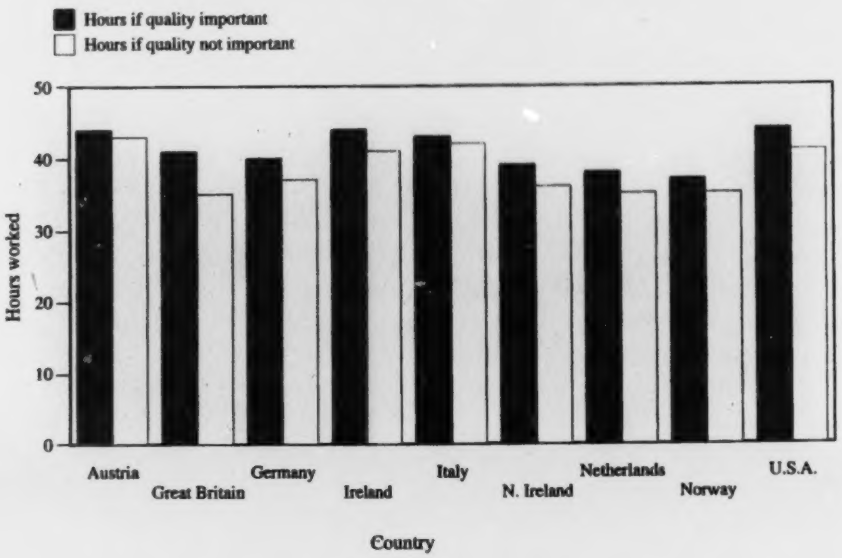
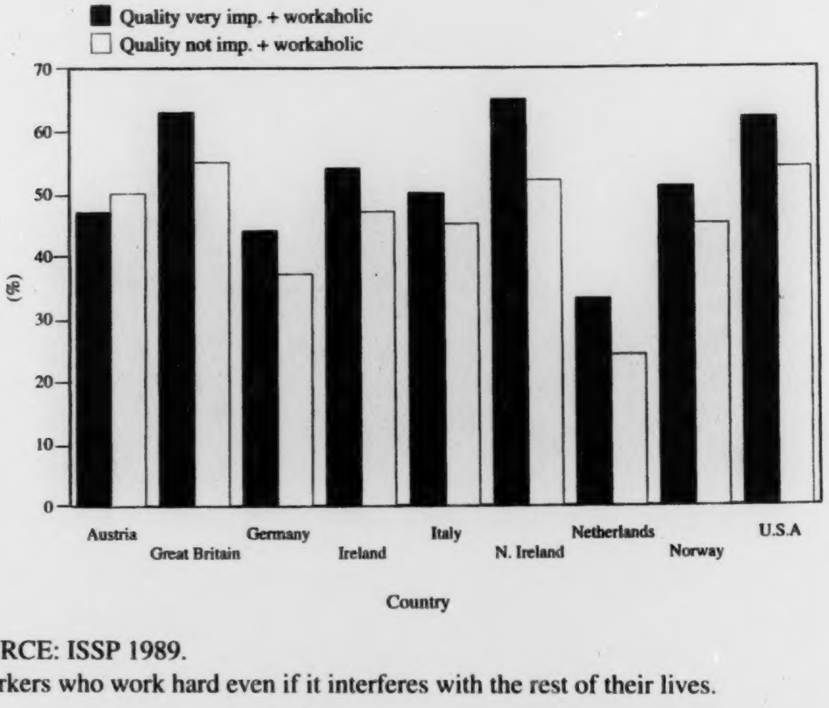


Figure 5B Workaholics[a] and Feelings about Importance of Work Quality for Pay



SOURCE: ISSP 1989.

[a] Workers who work hard even if it interferes with the rest of their lives.

Notes

1. The ISSP is a program of cross-national collaboration carried out with research institutes that conduct annual surveys of social attitudes and values. The virtue of the survey is that it seeks to ask similar questions in identical form in the participating nations.
2. In the ISSP there is a question about preferred hours of work that asks: Think of the number of hours you work and the money you earn on your main job, including overtime. Bell and Freeman (1995) analyze the responses to this question. It shows that Americans are more likely to want more hours and more pay than Europeans, but by much smaller amounts than are shown by questions that exclude overtime. The European Economic Community has asked workers about their preferences between increases in pay (for all hours of work) and reductions in hours worked that would maintain their real income in a collective bargaining session. The responses show that people prefer increases in pay, consistent with the hypothesis that most workers are in equilibrium in their choices about current hours of work.
3. Preferences for hours worked appears to be very closely linked to age in Japan, according to various work hour surveys (see Public Opinion Survey on Working Hours and Five-Day Workweek, for example), with younger Japanese disproportionately predisposed to shorter working hours than their more senior colleagues.
4. The problem is that with the same capital stock, fewer hours worked (due, say, to unemployment) implies higher labor productivity. Thus, Spain will have a relatively high productivity when employment falls—hardly an indicator of a good economic performance. All partial economic indicators are potentially misleading.
5. Let H = annual hours; F = hours of full-time workers; and p = proportion of workers who are part-time in the given country. Assume part-timers work $0.6F$ hours. Then $H = (1 - p)F + 0.6pF = (1 - 0.4p)F$. A change in p thus changes H by $0.4Fdp$. Dividing by H to get percentage changes in annual hours, we get $DH/H = 0.4/(1 - 0.4p)dp$.
6. Data from the 1989 ISSP suggest that at least with respect to perceptions, this may be true. U.S. workers (87 percent) are far more likely to indicate that "the quality of their work is important in determining their pay" than are German workers (47 percent).
7. There are two compelling reasons for concentrating our empirical analysis on full-time private nonagricultural workers. First, the theory that we build on to explain the wage-inequality link is based on a significant amount of future job attachment that is less likely to be an important component of the marginal decisions of part-time workers. Second, measurement error problems are likely to be exacerbated among workers reporting very low numbers of usual weekly hours.
8. Small numbers of observations in individual cells imply that estimates are sensitive to extreme values of hours or wages (either true or measured with error). Correlation estimates by major industry-occupation cells produced a qualitatively similar, although less strong relationship between hours worked and wage variation.

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Table A1 Mean Weekly Hours and Wages in Nonagricultural Industry, Full-Time Workers Only[a]

Industry	Weekly hours[b]	Std. dev. ln (hours)	Hourly wages	Std. dev. ln (hourly wage)
Mining	47.1	0.217	13.87	0.526
Transportation	45.1	0.180	10.71	0.484
Private household services	44.8	0.245	3.56	0.596
Petroleum and coal	44.6	0.134	12.64	0.300
Wholesale trade	44.1	0.140	11.10	0.487
Other professional services	43.8	0.170	12.75	0.623
Non-electrical machinery	43.5	0.122	12.71	0.467
Rubber and plastics	43.2	0.130	9.47	0.439
Repair service	43.2	0.140	8.75	0.436
Tobacco manufacturers	43.0	0.166	16.80	0.593
Motor vehicles and equipment	42.8	0.120	12.86	0.434
Food and kindred products	42.6	0.130	9.26	0.480
Business services	42.6	0.133	10.85	0.574
Stone, clay, and glass	42.6	0.097	11.40	0.410
Insurance and real estate	42.5	0.136	11.19	0.510
Fabricated metal	42.5	0.104	10.76	0.420
Chemicals and allied products	42.5	0.110	15.32	0.491
Education services	42.4	0.143	10.51	0.541
Electrical machinery	42.4	0.111	12.85	0.513
Retail trade	42.3	0.136	7.73	0.504
Leather and leather products	42.3	0.112	8.62	0.539
Entertainment and rec. services	42.3	0.136	8.61	0.518
Toys, amusements, and sporting	42.3	0.133	8.41	0.531
Aircraft and parts	42.3	0.104	12.97	0.458
Other transportation equipment	42.3	0.106	13.75	0.382
Primary metals	42.2	0.111	11.52	0.404
Utilities	42.2	0.113	14.00	0.425
Textile mill products	42.2	0.106	8.11	0.436
Construction	42.2	0.117	11.45	0.471
Paper and allied products	42.1	0.097	10.81	0.414
Communications	42.0	0.119	13.51	0.454

Industry	Weekly hours[b]	Std. dev. ln (hours)	Hourly wages	Std. dev. ln (hourly wage)
Professional and photo equipment	41.9	0.098	11.60	0.479
Lumber and wood, except furniture	41.9	0.096	8.40	0.339
Social services	41.8	0.150	8.41	0.555
Personal services excluding household	41.6	0.128	6.86	0.461
Printing and publishing	41.6	0.107	10.54	0.481
Hospitals	41.6	0.115	10.91	0.434
Banking and other finance	41.6	0.108	11.70	0.501
Furniture and fixtures	41.3	0.070	8.50	0.346
Miscellaneous manufacturing	41.2	0.078	7.92	0.509
Health services excluding hospitals	41.1	0.118	9.00	0.459
Apparel	40.2	0.057	6.56	0.448

[a] Full-time workers are defined to be workers with usual weekly hours ≥ 35.

[b] Usual weekly hours at main job.

Table A2 Mean Weekly Hours and Wages in Nonagricultural Detailed Occupation, Full-Time Workers Only[a]

Industry	Weekly hours[b]	Std. dev. ln (hours)	Hourly wages	Std. dev. ln (hourly wage)
Health diagnosis	57.0	0.197	16.79	0.527
Lawyers and judges	46.7	0.157	22.39	0.473
Motor vehicle operators	46.5	0.189	8.83	0.410
Supervisors, proprietors, and sales	46.3	0.165	11.25	0.507
Executives, administrators, and managers	45.2	0.154	14.73	0.546
Teachers, college	44.9	0.178	17.68	0.523
Other professional specialties	44.3	0.187	12.07	0.582
Sales reps, finance, and business services	44.2	0.148	12.36	0.562
Sales reps, commodities, excluding retail	44.1	0.133	13.79	0.516
Natural scientists	44.0	0.130	15.20	0.529
Private household services	43.8	0.211	3.25	0.586
Other transportation and material moving	43.8	0.166	10.51	0.431
Engineers	43.3	0.122	19.28	0.325
Mechanics and repairers	43.1	0.129	11.08	0.405
Construction laborers	43.1	0.134	8.81	0.445
Other precision production	43.0	0.129	11.41	0.455
Teachers, except college	43.0	0.144	8.94	0.570
Mathematical and computer scientists	43.0	0.117	17.74	0.372
Management	42.7	0.113	14.44	0.449
Supervisors–administrative support	42.5	0.122	11.63	0.432
Engineering and science technicians	42.1	0.100	12.62	0.415
Engineering and science	42.0	0.105	14.75	0.500

Industry	Weekly hours[b]	Std. dev. ln (hours)	Hourly wages	Std. dev. ln (hourly wage)
Fabricators, inspectors, and samplers	41.8	0.110	9.49	0.424
Construction trades	41.7	0.114	11.61	0.450
Handlers, equip. cleaners. laborers	41.7	0.114	7.38	0.432
Health assessment and treating	41.5	0.117	13.16	0.356
Machine operators, excluding precision	41.3	0.092	8.13	0.423
Food service	41.3	0.148	5.31	0.426
Personal services	41.2	0.118	7.18	0.467
Personal service occupations	41.1	0.118	6.49	0.518
Protective services	41.1	0.093	6.48	0.391
Sales related	41.0	0.034	11.19	0.151
Health technologists and technicians	40.9	0.093	10.14	0.329
Freight, stock, and material handlers	40.7	0.092	7.03	0.381
Health service	40.6	0.123	6.71	0.366
Computer equipment operators	40.6	0.075	9.33	0.420
Financial records, processing	40.4	0.070	7.96	0.318
Main and message distributing	40.4	0.041	7.87	0.430
Other administrative support, clerical	40.4	0.075	8.63	0.395
Cleaning and building services	40.3	0.078	6.46	0.346
Secretaries, stenographers, and typists	40.0	0.058	8.36	0.355

SOURCE: May 1989 CPS.

[a] Full-time workers are defined to be workers with usual weekly hours ≥ 35.

[b] Usual weekly hours at main job.

Table A3 Hours Regressions[a]

Dependent variables: ln (usual hours worked per week)

Independent variables						
ln (hourly wage)	0.147 (0.005)	0.101 (0.006)	0.085 (0.006)	0.059 (0.006)	0.141 (0.005)	0.080 (0.006)
ln (family income)	–0.007 (0.004)	0.004 (0.004)	0.003 (0.004)	–0.002 (0.003)	–0.002 (0.004)	0.002 (0.003)
Std. dev. of ln (hourly age)[b] (ind–occ cell)	0.062 (0.021)	0.044 (0.021)	0.076 (0.022)	–0.080 (0.024)	—	—
Std. dev. of ln (hourly wage)[c] (occ cell)	—	—	—	—	0.374 (0.033)	0.374 (0.035)
Marriage dummy	—	0.030 (0.006)	0.023 (0.006)	0.013 (0.006)	—	0.022 (0.006)
Female dummy	—	–0.121 (0.005)	–0.100 (0.006)	–0.089 (0.006)	—	–0.097 (0.006)
Less than high school	—	–0.012 (0.014)	–0.025 (0.014)	–0.008 (0.014)	—	–0.027 (0.014)
Some high school	—	–0.117 (0.013)	–0.119 (0.012)	–0.110 (0.012)	—	0.117 (0.012)
Some college	—	–0.094 (0.009)	–0.087 (0.009)	–0.085 (0.008)	—	–0.087 (0.008)
College graduate	—	0.017 (0.008)	0.026 (0.008)	0.018 (0.008)	—	0.020 (0.008)

College plus	—	0.015 (0.010)	0.040 (0.011)	0.020 (0.011)	—	0.026 (0.011)
Union dummy	—	0.005 (0.008)	–0.010 (0.008)	0.010 (0.008)	—	–0.007 (0.008)
Region dummy	No	Yes	Yes	Yes	No	Yes
Industry[d] dummy	No	No	Yes	No	No	Yes
Occupation[d] dummy	No	No	No	Yes	No	No
R^2	0.088	0.152	0.217	0.217	0.098	0.181
N	10,344	10,344	10,344	10,344	10,344	10,540

[a] For private, nonagricultural workers, including part-time workers, standard errors are in parentheses.

[b] Standard deviation in ln (hourly wage) in detailed industry/detailed occupation cell.

[c] Standard deviation in ln (hourly wage) in detailed occupation cell.

[d] Detailed occupation and industry dummy variables.

Part II

4

Working Time, Wages, and Earnings Inequality among Men and Women in Canada, 1981–1993

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The increasing polarization of employment earnings in Canada has been well documented (Myles, Picot, and Wannell 1988; The Economic Council of Canada 1991; Morissette, Myles, and Picot 1994; Burbidge, Magee, and Robb 1993; Beach and Slotsve 1996; Richardson 1994). Various dimensions of this rising polarization of earnings have also been explored, including the declining real and relative wages of young workers (Myles, Picot, and Wannell 1988; Davis 1992; Betcherman and Morissette 1994), and the relative stability of the education wage premium in Canada as compared to the United States. (Freeman and Needels 1991; Bar-Or et al. 1993; Morissette, Myles, and Picot 1994). In most of the work on rising earnings inequality, changes in the distribution of working time have been largely ignored. Yet the distribution of hours of work has changed significantly in the 1980s and early 1990s for a variety of reasons.

The shifts in labor demand often discussed in the earnings inequality literature could well be reflected through changes in hours of work, as well as through relative wages. Decreased demand for less-skilled workers due to the introduction of skill-biased technology, changes in trading patterns, or for any other reason could be reflected in declining relative wages for the less-skilled (a price response), declining relative hours of work (a quantity response), or both. Furthermore, if the demand shifts resulted in declining wages among the less-skilled, it could lead to a supply side response on the part of workers. They could withdraw some amount of labor at the lower wage rate, as has been suggested by Freeman (1994); Juhn, Murphy, and Topel (1991); and

Kuhn and Robb (1996). These possibilities would lead to declining hours of work among lower paid, less-skilled workers relative to the higher paid.

Such a possibility has been addressed in earlier work in Canada by Picot, Myles, and Wannell (1990); MacPhail (1993); Morissette, Myles, and Picot (1994); and Morissette (1995). They found that the polarization in hours worked did increase through the 1980s. More workers were working longer hours, more were working shorter hours, and fewer were working the number of hours one would expect to see in regular full-time jobs. Furthermore, they observed that the increasing polarization was such that it would tend to increase inequality in annual earnings; that is, the more highly paid were working relatively longer hours, and the less paid relatively shorter hours. There appears to be an association between a polarization in hours of work and that of annual earnings in Canada. This possibility has received less attention in the United States, where the rising inequality in employment earnings has if anything been greater than in Canada (Freeman 1994; Kuhn and Robb 1996). Earlier studies by Burtless (1990) and Moffitt (1990) concluded that the rise in earnings inequality in the United States are associated primarily with an increased dispersion in hourly wages, not hours of work.

The goal of this chapter is to extend this work in a number of ways. First, we update the earlier work, which focused exclusively on the 1980s, by using data from the new Survey of Labour and Income Dynamics (SLID) for 1993. This is the only data source in Canada with hourly wage rate data, a central variable for this analysis. Given the anecdotal stories related to the potential causes such as those listed above, the polarization in hours worked may be increasing in the 1990s. The second goal is to determine whether it is changes in weeks worked per year or hours per week that play the major role in the changing distribution of working time and its impact on earnings inequality. This has not been previously done.

The chapter finds that the rise in inequality among prime-age males and the decline among their female counterparts during the 1980s appears to be largely associated with changes in hours worked. Changes in the distribution of hourly wages played a much smaller role. This supports the earlier work that reached a similar conclusion. Changes in weeks worked (particularly for women) and, more importantly, in hours worked per week were both part of this phenomenon. Extending the analysis into the 1990s suggests that the same basic conclusion holds; the smaller observed change in annual earnings inequality is associated primarily with changes in hours of work, although changes in hourly wages do play a role. However, the rise in annual earnings inequality among males slowed over the 1984–1993 period compared with the 1981–1989 period. Annual earnings inequality among prime-age men and women *combined* is seen to change little over the past decade. Other work (OECD 1996; Zyblock 1996) also suggested that earnings inequality in Canada and a number of other countries stopped increasing during the 1990s. The exceptions are the United States and the United Kingdom, where rising inequality continues.

This chapter also asks to what extent changes in hours worked are associated with one of the more striking dimensions of earnings change in Canada, the declining real and relative earnings of young workers. Here, changing relative hours play a minor role.

Declining relative annual earnings among the young are associated with a decline in real and relative *hourly wages* among young males, and with both declining relative wages and declining relative weekly hours worked among young women. There has been a substantial downward adjustment in wage rates among younger workers in Canada during the 1980s and early 1990s.

THE DATA AND APPROACH

The Data

Two sets of data are employed. First, the Survey of Consumer Finances (SCF) for the period 1981 to 1993 is used to decompose the change in annual earnings inequality into that associated with changing weeks worked and changing weekly earnings. Ideally, one would use this source to assess the role of weekly hours worked and hourly wages as well, but such information is not available in the SCF for the reference year.

To further decompose changes in weekly earnings into that associated with hours worked per week and hourly wages, a series of special surveys are used. These are the Survey of Work History (SWH) for 1981, the Survey of Union Membership (SUM) for 1984, the Labour Market Activity Survey (LMAS) for 1986 and 1989, and the SLID for 1993. While the content of these surveys differs substantially, they all collect the variables in which we are primarily interested—hours worked per week and hourly wages—in a similar manner. More is said throughout the chapter regarding the similarities and differences among these surveys. The series of special surveys is the only source of comparable data over a reasonable period of time on hourly wages in Canada.

The Population

Since we are concerned with changes in hours worked and wages, we include workers who worked at any time during the year in the SCF sample. To maintain comparability with earlier studies, persons with self-employment earnings are excluded.[1] Other studies have restricted the population to full-time, full-year workers, but to do so would remove changes that have been taking place in weeks worked per year or in hours per week among much of the population. The analysis in the first section is restricted to prime-age workers (aged 25–54), however, to exclude possible events such as increasing part-time employment among students[2] and increasing early retirement rates among older workers. This chapter focuses on events occurring to workers in their prime working years.

As noted, we include only those persons who worked at some time during the year. Ideally the population should include everyone in the labor force, i.e., working or seeking work. Changes in hours of work should ideally include the increase or decrease in the number of persons working zero hours per year but wishing work. This is the group of workers unemployed all year, or unemployed for some portion of the year but not working during the year. These people have not been included in the sample because they do not have an observed wage rate. The significance of this omission depends upon the size of this group and whether it represents an increasing or declining share of the labor force.

In 1986, 1.4 percent of the male labor force were unemployed all year, and an additional 1.3 percent were unemployed part year but did not work. By 1989, these numbers had fallen to 0.7 percent and 0.9 percent as the economic expansion continued. In the recessionary period of 1993, they rose to 2.6 percent and 1.3 percent. Thus, the proportion of the labor force excluded from the sample that might have legitimately been included ranged from 1.8 percent to 3.9 percent, fairly small numbers. Most unemployed persons work at least part of a year, and hence are in our sample.[3]

The Time Frame

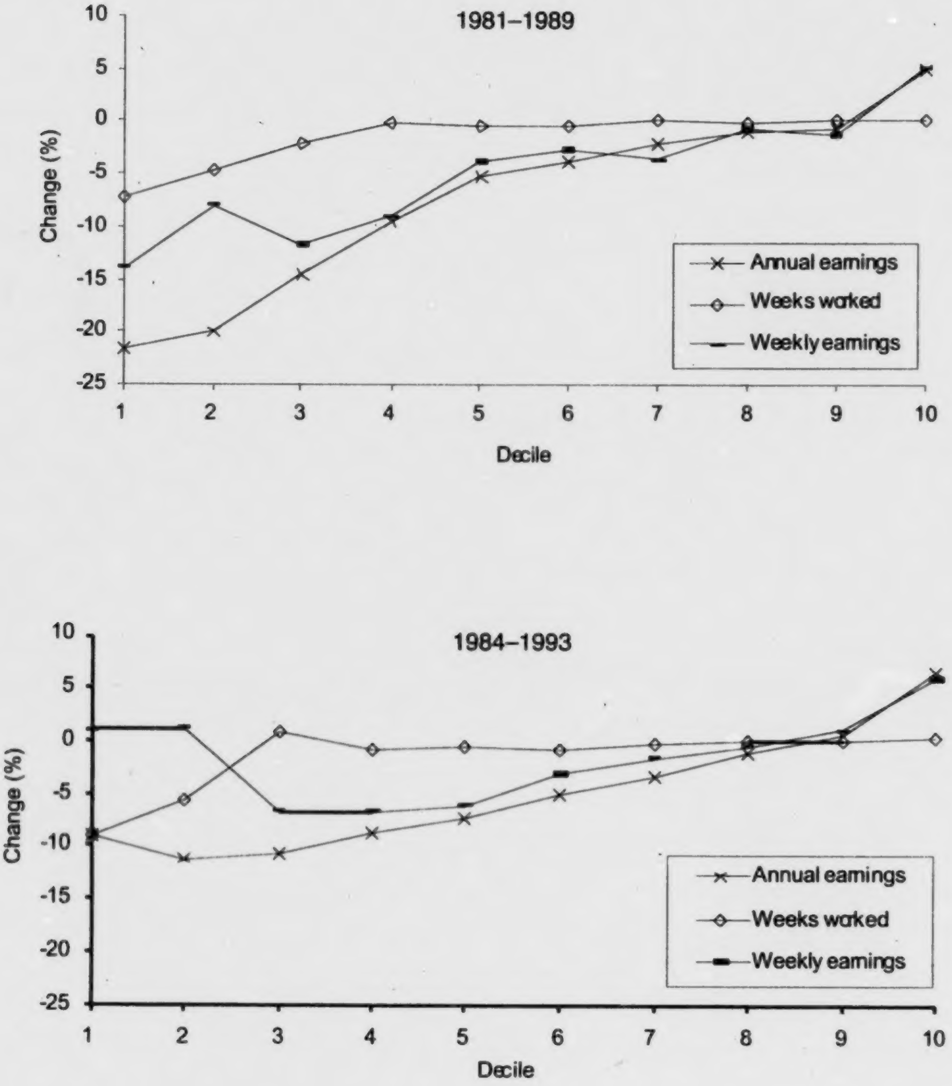
The dispersion of wages, hours, and earnings all vary significantly over the business cycle. Hence, years were selected that had comparable unemployment rates. The results are presented for two periods, 1981–1989 and 1984–1993. In terms of unemployment, these pairs of years are almost identical. Unemployment was at 7.6 percent and 7.5 percent respectively in the 1981–1989 period, and at 11.3 percent and 11.2 percent, respectively, in 1984 and 1993. Years prior to 1981 are excluded because data comparable to that from the special surveys are not available for the earlier years. However, other studies have shown that most of the increasing polarization in earnings occurred following 1981 (e.g., Morissette, Myles, and Picot 1994), so we are focusing on the period of most change.

HOURS, WAGES, AND INEQUALITY AMONG PRIME-AGE MALE WORKERS

The Rising Inequality of Annual Earnings: Associated with Changes in Weeks Worked per Year or Weekly Earnings?

The familiar increase in the earnings gap between the low and high earners is evident in the SCF data. Between 1981 and 1989, among 25- to 54-year-old prime-age males, real average annual earnings in the bottom three deciles fell 15 percent to 22 percent while rising 5 percent in the top decile (Figure 1). The gap in real earnings between high and low prime-age male earners opened considerably over this period.

Figure 1 Percent Change in Annual Earnings, Weeks Worked, and Weekly Earnings, by Decile,[a] Men, 25–54



SOURCE: Survey of Consumer Finances.

[a] Deciles based on annual earnings; in Canadian dollars.

The majority of the change in earnings was associated with changes in weekly earnings, less was related to changes in weeks worked. This can be seen in Figure 1, as the percentage change in annual earnings for different deciles between 1981 and 1989 largely reflects the change in weekly earnings. For example, workers at the bottom of the annual earnings distribution worked an average of 31 weeks per year in 1981 and 28.7 weeks by 1989, a decrease of 7 percent. However, weekly earnings among this lowest earning group fell fully 14 percent. The decline in these two variables resulted in the overall drop of 22 percent in real annual earnings. The number of weeks worked changed little in the middle and upper deciles; all of the change was in weekly earnings.[4]

Thus, among prime-age males during the 1980s, both changes in weeks worked and weekly earnings were associated with the rising gap between the bottom and the top of the earnings distribution, although changes in weekly earnings tended to dominate.

An alternative means of demonstrating this is to turn to a summary measure of the degree of inequality or dispersion in a distribution, the variance of the logarithm. The advantage of this measure is that it is decomposable. Annual earnings is a product of weeks worked and weekly earnings. The total dispersion in the distribution of annual earnings can be decomposed into that due to the dispersion in weeks worked, that due to the inequality in weekly earnings, and a term indicating the covariance between these two factors.

Among prime-age male earners, the variance of the log of annual earnings rose almost 30 percent between 1981 and 1989 (Table 1)—a very significant increase in inequality. Fifty-three percent of this rise was due to the increased dispersion in weekly earnings, and another 23 percent to the increasing covariance between weekly earnings and weeks worked. The tendency for full-year workers to have higher weekly earnings and part-year workers to have lower weekly earnings increased during the 1980s. Only 23 percent of the rise in earnings inequality was due to an increasing dispersion of weeks worked.

Thus, over the 1980s, for prime-age males, the considerable increase in polarization in earnings was largely associated with changes in weekly earnings rather than weeks worked per year.

Table 1 Change in Variance of the Log of Annual Earnings, Weekly Earnings, and Weeks Worked, Prime-Age Men[a]

					Change	
Variable	1981	1984	1989	1993	1981–89	1984–93
Annual earnings	0.468	0.702	0.605	0.771	0.137	0.069
Weeks worked	0.119	0.199	0.151	0.249	0.032 (23%)[b]	0.050 (72%)
Weekly earnings	0.297	0.379	0.370	0.400	0.073 (53%)	0.021 (30%)
2 × covariance	0.052	0.124	0.084	0.122	0.032 (23%)	–0.002 (–3%)

SOURCE: Survey of Consumer Finances.

[a] In Canadian dollars.

[b] The number in parentheses indicates the percent distribution.

The 1984–1993 period tells a somewhat different story.[5] Inequality in annual earnings increased relatively little among prime-age males. The variance of the log rose by about 10 percent (0.069 points), compared to the 30 percent increase between 1981 and 1989, the business cycle peaks. The major rise in earnings inequality, and the associated decline in earnings of lower paid workers (see Figure 1), occurred largely in the very early 1980s. Other evidence suggests that the upward trend in earnings inequality may have ceased in the 1990s, at least among full-time full-year workers (OECD 1996; Zyblock 1996).

The slower rise in annual earnings inequality between 1984 and 1993 is associated with a slowdown in the growth of inequality in *weekly* earnings. Inequality in this variable rose only 6 percent, compared to 25 percent over the 1981–1989 period. The rise in inequality in weeks worked was about the same in the two periods, roughly 25 percent as measured by the variance of the logs. Thus, changes in inequality in weekly earnings played a dominant role in annual earnings inequality among prime-age males.

The Rising Polarization in Weekly Earnings: Associated with Changes in Hours per Week or Hourly Wages?

Weekly earnings can be thought of as a function of two factors: hours worked per week and hourly wage rate. To try to decompose the observed changes in weekly earnings, we turn to new data sources. The Survey of Consumer Finances, used in the previous section, does not include data on hours per week or hourly wages for the reference year. Thus we turn to a series of special data sources[6] that collected information on wages and hours worked per week. Weekly earnings are computed for these data sets simply by multiplying hourly wages by hours per week. The hours and earnings data refer to the major job held by the worker in December, as this was the only information available on the 1984 survey.[7]

This same question—"Is the rise in weekly earnings inequality associated with changes in hours per week or hourly wages?"—was addressed by Morissette, Myles, and Picot (1994) and Morissette (1995) for the 1980s. Both studies concluded that most of the change in weekly earnings during that decade was associated with changes in hours per week, and increasing covariance between hours worked and wages. We revisit that question and extend the analysis to the early 1990s.

First, hours worked. There was a substantial increase in the polarization of hours worked per week over the 1980s. Labour Force Survey (LFS) data suggest that between 1981 and 1989, for example, the proportion of men working 35–40 hour weeks in their main jobs fell from 77 percent to about 73 percent and continued to fall to around 69 percent by 1993.[8] The proportion working longer hours (more than 40) increased from 18 percent in 1981 to 22 percent in 1989, and 24 percent in 1993.[9] The proportion working shorter hours (under 35) also rose, from 4 percent in 1981 to 5 percent in 1989, reaching 8 percent by 1993. Thus, fewer men were working regular 35–40 hour weeks, and more were working both shorter and longer hours. These trends are presented in more detail in Morissette and Sunter (1994).

The distribution of prime-age male workers by hourly wage,[10] the other major determinant of weekly earnings, presents a very different story. There has been relatively little change in this distribution over the 1980s and early 1990s. The distribution of workers by their hourly wage, in constant 1993 dollars, is shown in Table 2. The change in the distribution between 1981 and 1989 shows no systematic shift toward the top or bottom. The same is observed for the 1984–1993 period.

The variance of the log is used to decompose the rise in the inequality in weekly earnings (Table 3). The results for prime-age males over the 1980s suggests that the majority of the rise in inequality in weekly earnings was associated with the rise in inequality in hours

Table 2 Distribution of Prime-Age Men, by Hourly Wage in Their Jobs[a] (%)

Hourly wage[b] ($)	1981	1984	1989	1993	Change 1981–89	Change 1984–93
0–7.00	5.1	3.2	3.6	4.5	–1.5	+1.3
7.01–8.80	5.2	4.3	4.3	4.4	–0.9	+0.1
8.81–10.80	8.2	6.6	6.9	5.8	–1.3	–0.8
10.81–12.90	8.9	9.5	10.2	9.2	+1.3	–0.3
12.91–14.90	10.7	12.1	12.1	10.0	+1.4	–2.1
14.91–16.90	12.3	9.8	11.0	10.6	–1.3	+0.8
16.91–19.10	12.9	13.6	12.3	12.3	–0.6	–1.3
19.11–21.90	12.3	14.4	12.7	15.4	+0.4	+1.0
21.91–27.60	14.6	17.2	16.0	16.4	+1.4	–0.8
> 27.60	9.7	9.3	10.8	11.6	+1.1	+2.3

SOURCE: Survey of Work History (1981); Survey of Union Membership (1984); Labour Market Activity Survey (1986, 1989); Survey of Labour and Income Dynamics (1993).

[a] Major job held in December.

[b] In constant 1993 Canadian dollars.

Table 3 Change in the Variance of the Log of Weekly Earnings, Hours Per Week, and Hourly Wages, Prime-Age Men[a]

Variable	1981	1984	1989	1993	Change 1981–89	Change 1984–93
Weekly earnings	0.250	0.258	0.287	0.310	0.037	0.052
Hours per week	0.054	0.064	0.086	0.099	0.032	0.035
Hourly wage	0.222	0.194	0.214	0.198	-0.008	0.004
2 × covariance	–0.026	0.000	–0.012	0.013	0.014	0.013

SOURCE: Survey of Work History (1981); Survey of Union Membership (1984); Labour Market Activity Survey (1986, 1989); Survey of Labour and Income Dynamics (1993).

[a] In Canadian dollars.

worked per week (0.032 of a total increase of 0.37 in the variance of the log), while hourly wages played little or no role.[11] The findings are similar for the 1984–1993 period. Among prime-age males, weekly earnings inequality rose 20 percent, and of this increase, two-thirds was related to changes in the distribution of hours worked per week; hourly wages played only a minor role.

Thus, the data from the special surveys suggest that inequality in hourly wages has risen only marginally when comparable points in the business cycle are used. Changes in the inequality of weekly earnings appear to be largely associated with changes in hours worked per week. This result was observed in earlier studies for the 1980s, and it would seem that it holds for the 1984–1993 period as well.[12,13]

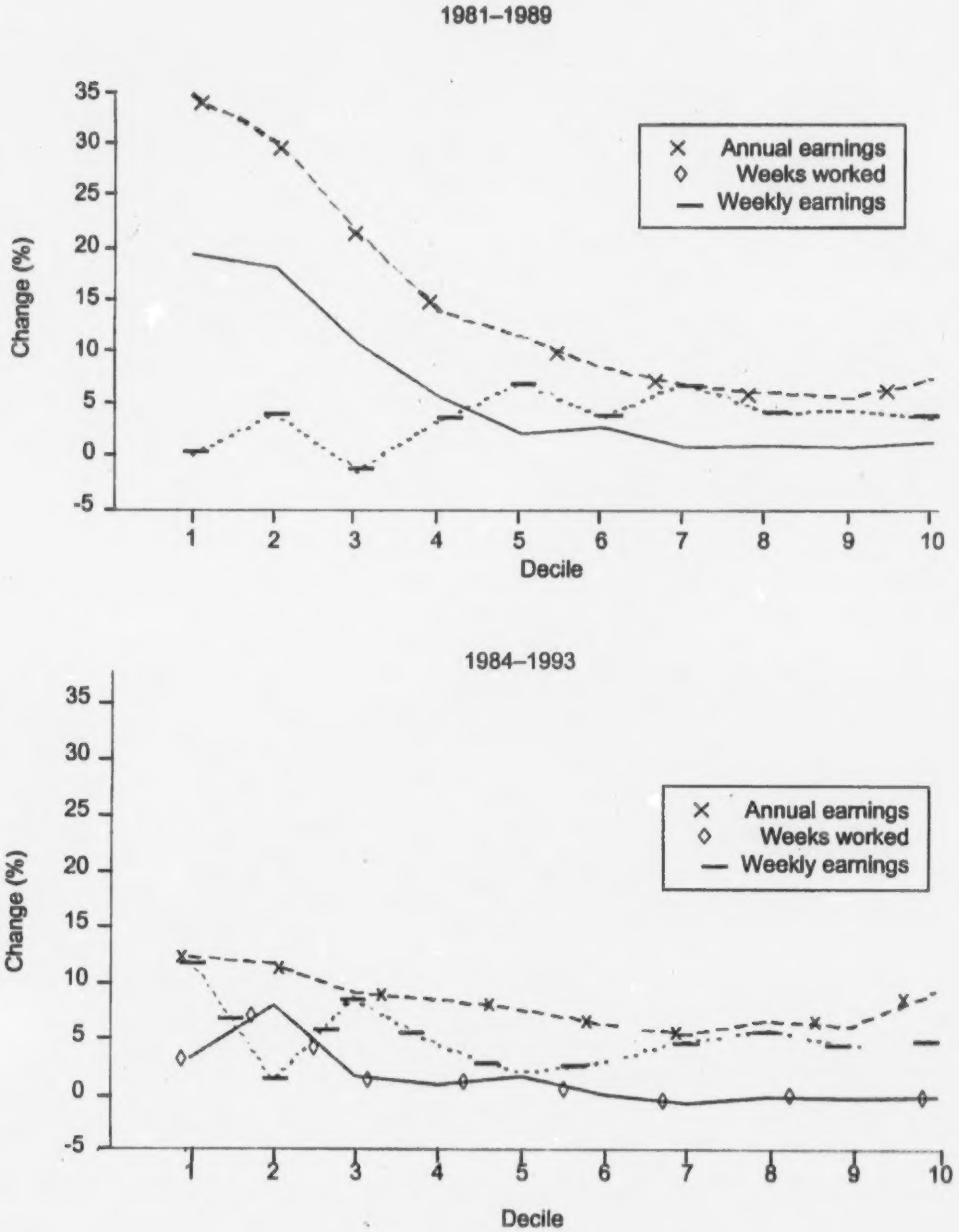
HOURS, WAGES, AND INEQUALITY AMONG PRIME-AGE FEMALE WORKERS

Polarization of Annual Earnings: Associated with Changes in Weeks Worked per Year or Weekly Earnings?

Not surprisingly, the trends are very different for men than for women, since their underlying employment and earnings trends are very different. The earnings gap between men and women has been closing in Canada, and an increasing share of women are working, whereas among men the share working has been falling.[14]

Annual earnings have not fallen among women at the bottom of the earnings distribution as they have among men. In fact, they have increased quite dramatically. Between 1981 and 1989, annual earnings among the lowest paid women rose 35 percent, largely because they were working more weeks per year in 1989 than in 1981. Weeks worked rose 19 percent among this group, whereas average weekly earnings did not change at all (Figure 2).

Figure 2 Percent Change in Annual Earnings, Weeks Worked, and Weekly Earnings, by Decile,[a] Women, 25–54



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[a] Deciles based on annual earnings, in Canadian dollars.

With such a large rise in earnings among low paid women, not surprisingly annual earnings inequality declined among prime-age females during both the 1981–1989 period (declining 0.154 as measured by the variance of the log); and the 1984–1993 period (declining 0.037) (Table 4). In both cases this was associated with a decline in the inequality of weeks worked (i.e., lower paid women working relatively more weeks per year). During 1981–1989, two thirds of the decline in annual earnings inequality was associated with a decline in the inequality of weeks worked; during 1984–1993 it was 59 percent. There was little change in the inequality of weekly earnings. This is in contrast to men, where there was relatively little change in weeks worked, but a rise in inequality of weekly earnings.

Changes in Hours Worked per Week and Hourly Wages among Women

Like men, a smaller share of women are working a regular 35–40 hour week. According to the Labour Force Survey, the proportion of women working 35–40 hours per week fell from 68 percent in 1981 to 65 percent in 1989, and 61 percent in 1993 (Morissette and Sunter 1994).[15] There are slightly more working fewer hours (the share working under 20 rose from 20 percent in 1981 to 22 percent in 1993), and more working longer hours (over 40 hours rose from 5 percent to 8 percent). Thus, there has been an increase in the polarization of hours

Table 4 Change in Variance of the Log of Annual Earnings, Weekly Earnings, and Weeks Worked, Prime-Age Women[a,b]

					Change	
	1981	1984	1989	1993	1981–89	1984–93
Women 25–54						
Annual earnings	1.077	1.057	0.923	1.020	–0.154	–0.037
Weeks worked	0.362	0.340	0.262	0.319	–0.100 (65%)	–0.022 (59%)
Weekly earnings	0.571	0.552	0.530	0.552	–0.041 (26%)	0.000 (0%)
2 × covariance	0.144	0.164	0.132	0.150	–0.012 (8%)	–0.015 (41%)

SOURCE: The Survey of Consumer Finances.

[a] In Canadian dollars.

[b] The numbers in parentheses indicates the percent distribution.

worked among women, just as there has been among men. Such a polarization in hours worked does not necessarily lead to increased polarization of earnings, however, since it depends upon whether high or low wage earners are increasing their hours of work.

Unlike for men, the hourly wage distribution has changed quite significantly among prime-age women (Table 5). A larger share of women are earning higher hourly wages. For example, the proportion of women earning over $17.00 per hour rose from 24.6 percent in 1984 to 31.2 percent in 1993. Among men there was virtually no change, although a larger share of men than women still earned over $17.00 per hour in 1993 (55.7 percent against 31.2 percent). Among women there was a corresponding decline in the proportion earning low hourly wages.

Table 5 Distribution of Prime-Age Women, by Hourly Wage in Their Jobs[a] (%)

Hourly wage[b] ($)	1981	1984	1989	1993	Change 1981–89	Change 1984–93
$ 0–7.00	14.0	11.6	11.8	11.3	–2.2	–0.3
7.01–8.80	12.1	12.9	10.5	8.6	–1.6	–4.3
8.81–10.80	14.6	13.3	13.3	10.7	–1.3	–2.6
10.81–12.90	14.2	15.9	16.7	14.3	+2.5	–1.6
12.91–14.90	12.7	13.5	13.3	13.1	+0.6	–0.4
14.91–16.90	9.0	8.3	8.7	10.9	–0.3	+2.6
16.91–19.10	7.1	7.9	7.9	8.0	+0.8	+0.1
19.11–21.90	6.3	7.4	7.0	9.0	+0.7	+1.6
21.91–27.60	6.0	6.5	6.4	8.8	+0.4	+2.3
> 27.60	3.9	2.8	4.3	5.4	+0.4	+2.6

SOURCE: Survey of Work History (1981); Survey of Union Membership (1984); Labour Market Activity Survey (1986, 1989); Survey of Labour and Income Dynamics (1993).

[a] Major job held in December.

[b] In constant 1993 Canadian dollars.

The decomposition of the variance of the log displays the association between these changes and weekly earnings inequality among prime-age women (Table 6). Overall, the story is one of little change. Weekly earnings inequality among prime-age women has changed little, as noted earlier, particularly during the 1981–1989 period. During the 1984–1993 period, where inequality does rise marginally according to the special surveys data (but with no change registered in the SCF data), this is largely driven by the changing distribution of hours worked. Thus, overall there is little change in weekly earnings inequality, and where it is observed, it is related to changes in hours worked.

Inequality in Earnings, Hours, and Wage for All Prime-Age Workers

Different trends among men and women begs the question of whether the labor market has been redistributing earnings to prime-age workers as a whole in a fundamentally different manner in the late 1980s and early 1990s compared to earlier periods. And if so, what is the role of changes in working time and hourly wages.

Based on the variance of the log, annual earnings inequality among prime-age men and women combined has changed relatively little over the 1980s or early 1990s (Table 7). If anything, these data suggest that

Table 6 Change in the Variance of Log of Weekly Earnings, Hours Per Week, and Hourly Wages, Prime-Age Women[a]

					Change	
Variable	1981	1984	1989	1993	1981–89	1984–93
Weekly earnings	0.470	0.471	0.472	0.526	0.002	0.055
Hours per week	0.229	0.224	0.206	0.290	–0.023	0.066
Hourly wage	0.250	0.228	0.241	0.220	–0.009	–0.008
2 × covariance	–0.009	0.020	0.026	0.016	0.035	–0.004

SOURCE: Survey of Work History (1981); Survey of Union Membership (1984); Labour Market Activity Survey (1986, 1989); Survey of Labour and Income Dynamics (1993).

[a] In Canadian dollars.

Table 7 Change in the Variance of the Log of Annual Earnings, Weekly Earnings and Weeks Worked, All Prime-Age Workers[a]

					Change	
Variable	1981	1984	1989	1993	1981–89	1984–93
Annual earnings	0.910	0.969	0.855	0.956	–0.055	–0.013
Weeks worked	0.233	0.264	0.206	0.283	–0.027	0.019
Weekly earnings	0.529	0.540	0.523	0.531	–0.006	–0.009
2 × covariance	0.148	0.165	0.125	0.142	–0.023	–0.023

SOURCE: Survey of Consumer Finances.

[a] In Canadian dollars.

earnings inequality declined slightly between 1981 and 1989 (6 percent), as well as between 1984 and 1993 (1 percent). These are, however, relatively small changes.

As noted before, the two sets of data, SCF and the special surveys, provide somewhat different results regarding the change in inequality in *weekly* earning. There is little change in the SCF, and a small increase in the special surveys (from 7 percent to 10 percent). In the special surveys this increase is due to changes in the inequality in hours worked per week, or the increasing covariance between hours and wages.[16] The inequality in hourly wages changes little over either period for prime-age workers (Table 8).

Table 8 Change in the Variance of the Log of Weekly Earnings, Hours Per Week, and Hourly Wages, All Prime-Age Workers[a]

					Change	
Variable	1981	1984	1989	1993	1981–89	1984–93
Weekly earnings	0.416	0.432	0.444	0.478	+0.028	+0.046
Hours per week	0.144	0.149	0.156	0.209	+0.012	+0.060
Hourly wage	0.254	0.235	0.248	0.223	–0.006	–0.012
2 × covariance	0.017	0.048	0.040	0.046	+0.023	–0.002

SOURCE: Survey of Work History (1981); Survey of Union Membership (1984); Labour Market Activity Survey (1986, 1989); Survey of Labour and Income Dynamics (1993).

[a] In Canadian dollars.

Given this relative stability in inequality among prime-age workers, what has happened to the real earnings of workers earning lower wages? Basically, the annual earnings of prime-age workers at the bottom of the distribution have risen in both periods because of the tendency of women to work more weeks per year during the 1980s. There have been significant declines in the annual earnings of workers in the middle of the distribution over both periods, with small increases at the top.

SUMMARY OF THE ROLES OF HOURS WORKED AND WAGES IN EARNINGS INEQUALITY

Overall, most of the significant changes in earnings inequality among both men and women were associated with changes in working time, either weeks worked per year or hours per week, rather than changes in hourly wages. More specifically:

1) Inequality in annual earnings among prime-age males rose significantly between 1981–1989 (29 percent), and less so over 1984–1993 (10 percent). Other work (OECD 1996) indicates that among full-time, full-year workers annual earnings inequality fell slightly during the 1990s in Canada and numerous other countries, after having increased substantially in the 1980s. Annual earnings inequality among prime-age women fell during both periods.[17]

2) During both periods, the increase in male annual earnings inequality was associated with rising inequality in hours worked per week and weeks per year. Little of the rise was associated with rising inequality or polarization in hourly wages. Similarly, during both periods the decline in earnings inequality among women was largely associated with increasing weeks worked among women with low annual incomes.

3) Prime-age male workers at the bottom end (bottom three deciles) of the earnings distribution saw their annual earnings fall dramatically over the 1981–1989 period. This was associated with a decline in both weeks worked per year, and weekly earnings. The

change in the latter was largely associated with changes in hours worked per week. Prime-age women at the bottom of the earnings distribution saw their annual earnings rise, largely because they were working more weeks per year.

During the 1984–1993 period annual earnings fell among male workers in the bottom three deciles, but not so dramatically. This was largely due to a decline in weeks worked, as weekly earnings fell only marginally. Women at the bottom of the distribution saw their annual earnings rise as a result of working more weeks per year, but again the increase was not as dramatic as during the earlier period.

4) Both prime-age men and women at the top of the earnings distribution (top three deciles) saw their annual earnings rise during both periods due to an increase in weekly earnings, which was in turn associated with a substantial increase in hours worked per week, and some increase in hourly wages.

5) Among all prime-age workers (men and women combined), inequality in annual earnings, weeks worked, weekly earnings, and hourly wages has been quite stable over both periods, and where there have been changes, they are relatively small. Inequality in annual earnings and hourly wages has changed little.

Thus, changes in working time, primarily hours per week but also weeks per year, have been associated with the rise in earnings inequality among prime-age males and a decline among prime-age women[18] over both periods, although the change was much larger in the first than the second period. The best available data on wages (from the special surveys) suggest that while changes in hourly wages have played some role in the rise in earnings inequality, it is in no way dominant.

THE RISING EARNINGS DIFFERENTIAL BETWEEN THE YOUNG AND THE OLD—IS IT WAGES- OR HOURS-BASED?

A major dimension of the rising inequality story has been the increasing gap in annual earnings between younger and older workers, particularly among men. This has been largely due to declines in real earnings among young workers and has been well documented in Canada (Myles, Picot, and Wannell 1988; Betcherman and Morissette 1994; Morissette, Myles, and Picot 1994; Beaudry and Green 1996), and for other industrialized countries (Davis 1992). Our goal is to determine the extent to which this rising gap is associated with changes in hourly wages on one hand, or hours worked (either changes in weeks per year or hours per week) on the other. We focus on all workers, including part-time and full-time. This is necessary if one is to capture all the changes taking place in patterns of working time.

Changing Annual Earnings among Young Males

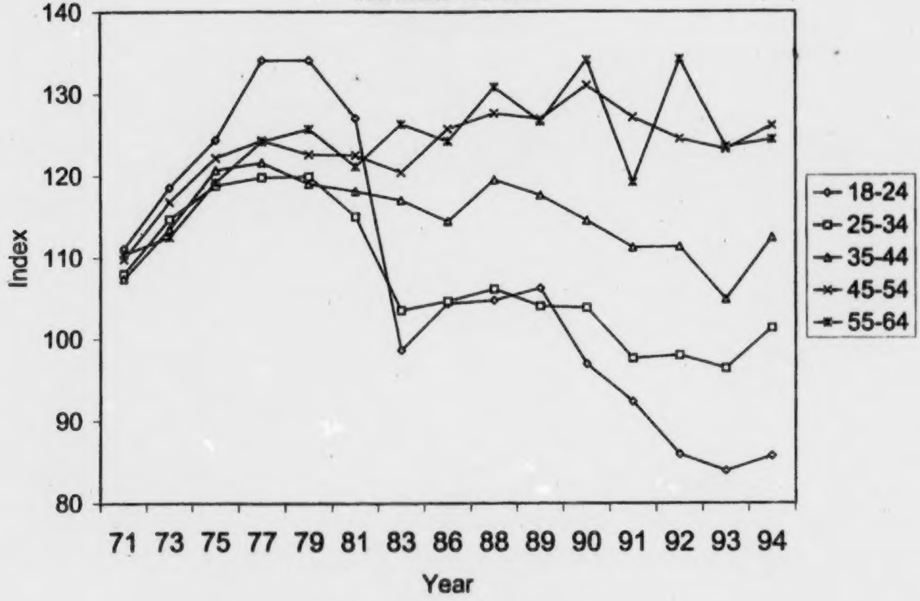
The much discussed decline in real average annual earnings among very young males (aged 18–24) occurred largely in the early 1980s, between 1981 and 1983 (Figure 3). Of the 36 percent drop in earnings between 1979 and the early 1990s, over 70 percent of it occurred in and around the 1981–1982 recession. There was little recovery during the late 1980s, and a subsequent smaller decline during the recession of the 1990s. Among 25- to 34-year-olds, whose earnings have fallen about 14 percent since 1979, most of the decline was in the 1980s recession, again with no recovery through the 1980s. These results included all young workers, full- and part-time. There have been significant changes in the number of young people working part-time over this period. However, the same general pattern is observed for full-time, full-year workers (Figure 3).

The approach here is to simply examine the relative values of four variables: annual earnings, annual weeks worked, hours per week, and hourly wage rate. The ratios computed are the average value of the variable for younger workers (under 35) relative to those for older workers (over 35). The change in these relative values over time will inform us as to which variables are associated with the decline in relative annual earnings among young male workers.

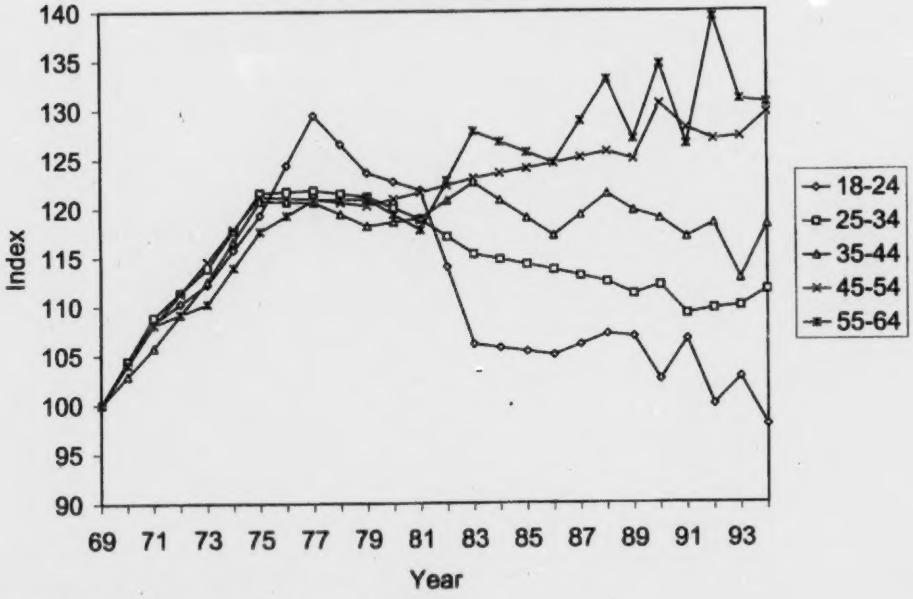
For men, the ratio of the average annual earnings of workers under 35 to those of workers over 35 fell from 0.64 in 1981 to 0.57 in 1989, and 0.55 in 1993 (Figure 4). Was this decline associated with a fall in relative weeks worked, hours per week, or hourly wages? The answer appears to be hourly wages. Relative weeks worked changed little over

Figure 3 Indexed Real Annual Wages and Salaries of Male Earners, 1969–1994 (1969=100)

All male earners

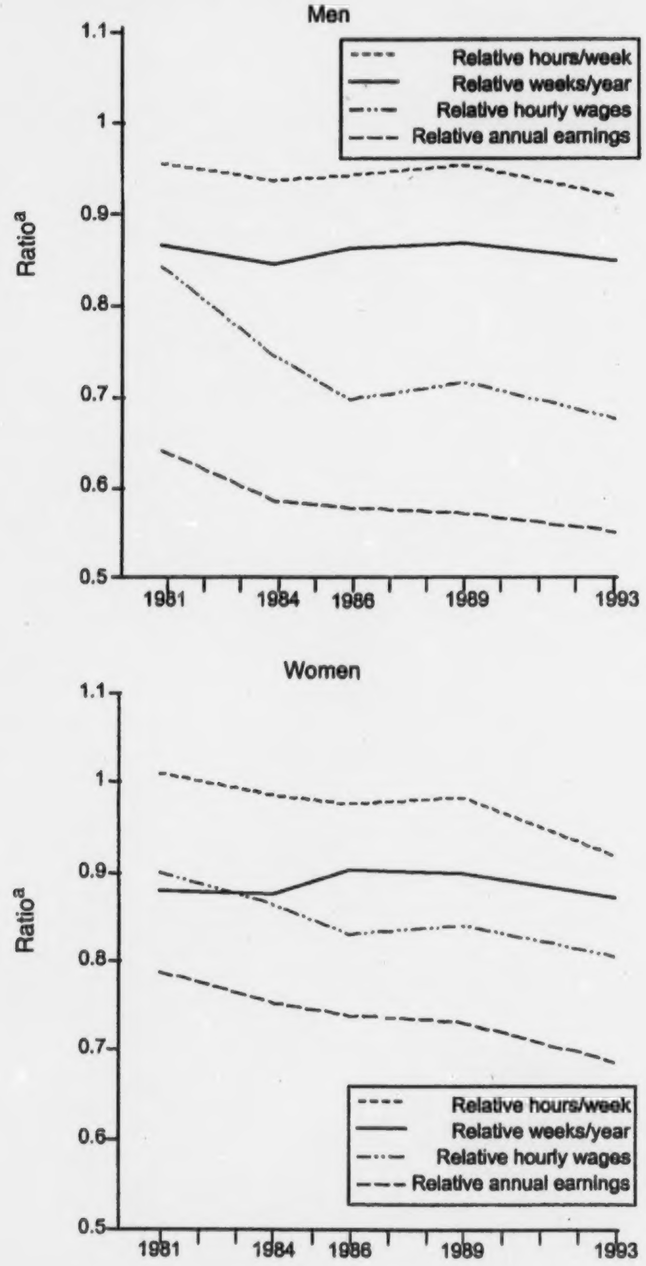


Full-year full-time male workers



SOURCE: Survey of Consumer Finances.

Figure 4 Relative Measures of Earnings and Hours: Workers under 35 Relative to Those over 35



[a] For example, for annual earnings this is the ratio of average annual earnings among 17- to 34-year-olds to the earnings of 35- to 64-year-olds.

the period (at around 0.87), as did relative hours worked per week (at 0.95). Relative hourly wages, however, fell from 0.84 in 1981 to 0.71 in 1989, and to 0.67 in 1993. This would suggest that virtually all of the decline in relative annual earnings among the young between 1981 and 1989 was associated with falling relative hourly wages.

To ensure that this result is not specific to this particular way of grouping younger and older workers, this same approach is used for other combinations of workers, and results compared for two periods, 1981–1989 and 1984–1993. Two separate younger age groups, 17–24 and 25–35, are compared with the central age group of older workers, those 45–54 (Table 9). Generally speaking, changing relative wages dominate the other variables.[19] Overall, however, the decline in relative (and real) annual earnings among the young appears to be associated with changing relative hourly wages, not changing relative hours worked per year.

Changing Annual Earnings among Young Women

As with the men, the real annual earnings of young women fell between 1977 and 1983 (about 21 percent), recovered somewhat during the 1980s, and fell again in the 1990s recession and beyond. The fall between the late 1970s and the mid 1990s was 29 percent. This was for all workers. Among women working full-time, full-year, the story was one of a decline in the early 1980s, followed by recovery in the late 1980s and early 1990s, so that there was little change overall during the entire period (Figure 5). Average earnings among older women (over 24) rose in all age categories, but more so among the older than younger workers, so that the relative annual earnings of 25-34-year-olds fell from 0.8 in 1981 to 0.7 in 1993. The relative decline was greater among full-time, full-year workers.

Were these declines in the relative earnings associated with declining relative wages or working time? The answer is both. Relative hourly wages have fallen among women under 35 (relative to those over 35), from 0.9 in 1981 to 0.8 in 1993. But relative hours worked per week have fallen as well, from 1 to 0.9. Thus, both have contributed to the decline in relative annual earnings. This is true for other populations as well (Table 10).

Table 9 Earnings and Hours Ratios of Younger Relative to Older Male Workers, Selected Years[a]

Age groups	Annual earnings[b]	Weeks worked[b]	Hours per week[c]	Hourly wages[c]
17–34 relative to 35–64				
1981	0.638	0.865	0.955	0.843
1984	0.584	0.845	0.937	0.743
1989	0.571	0.867	0.954	0.713
1993	0.550	0.849	0.919	0.673
Change 1981–89	–0.067	0.002	–0.001	–0.130
Change 1984–93	–0.034	0.004	–0.018	–0.070
17–24 relative to 45–54				
1981	0.392	0.743	0.871	0.703
1984	0.320	0.711	0.840	0.540
1989	0.316	0.731	0.840	0.508
1993	0.255	0.697	0.742	0.441
Change 1981–89	–0.076	–0.012	–0.031	–0.195
Change 1984–93	–0.065	–0.014	–0.098	–0.099
25–34 relative to 45–54				
1981	0.827	0.971	01.003	0.966
1984	0.764	0.944	0.992	0.851
1989	0.712	0.942	1.016	0.812
1993	0.682	0.935	0.987	0.741
Change 1981–89	–0.115	–0.029	+0.013	–0.154
Change 1984–93	–0.082	–0.009	–0.005	–0.110

[a] In Canadian dollars.

[b] Survey of Consumer Finances.

[c] Survey of Work History (1981); Survey of Union Membership (1984); Labour Market Activity Survey (1986, 1989); Survey of Labour and Income Dynamics (1993).

Table 10 Earnings and Hours Ratios of Younger Relative to Older Female Workers,[a] Selected Years

Age groups	Annual earnings[b]	Weeks worked[b]	Hours per week[c]	Hourly wages[c]
17–34 Relative to 35–64				
1981	0.787	0.878	1.009	0.900
1984	0.751	0.874	0.983	0.863
1989	0.729	0.898	0.982	0.838
1993	0.686	0.870	0.917	0.803
Change 1981–89	–0.058	0.020	–0.027	–0.062
Change 1984–93	–0.065	–0.004	–0.066	–0.060
17–24 Relative to 45–54				
1981	0.621	0.802	0.972	0.815
1984	0.540	0.780	0.919	0.700
1989	0.492	0.780	0.899	0.662
1993	0.382	0.741	0.759	0.572
Change 1981–89	–0.129	–0.022	–0.073	–0.153
Change 1984–93	–.0158	–0.039	–0.016	–0.128
25–34 Relative to 45–54				
1981	1.009	0.947	1.028	1.069
1984	0.983	0.949	1.029	1.042
1989	0.902	0.959	1.038	0.997
1993	0.836	0.933	0.979	0.909
Change 1981–89	–0.107	0.012	0.010	–0.072
Change 1984–93	–0.147	–0.016	–0.050	–0.133

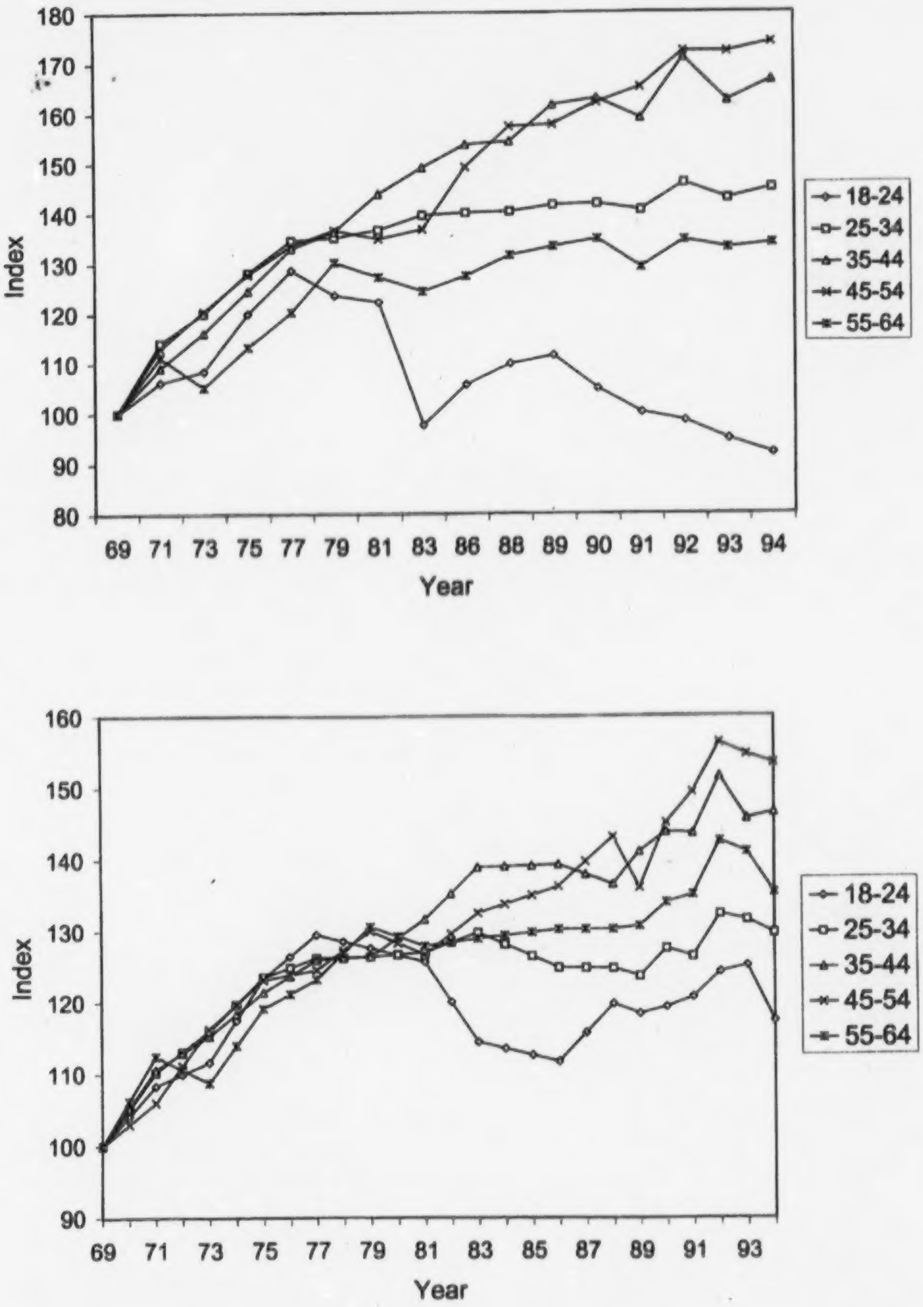
[a] In Canadian dollars.

[b] Survey of Consumer Finances.

[c] Survey of Work History (1981); Survey of Union Membership (1984); Labour Market Activity Survey (1986, 1989); Survey of Labour and Income Dynamics (1993).

Overall, the decline in annual earnings of the young appear to have more to do with changes in relative wages than in working time. This is unlike the story for changing inequality as a whole, where changes in working time play a major role. We address possible reasons for this in the concluding remarks.

Figure 5 Indexed Real Annual Wages and Salaries of Female Workers, 1969–1994 (1969=100)



DISCUSSION

The Association between Hours and Earnings

For hourly paid workers, the association between hours worked and annual earnings is direct; one works an extra hour and one's annual earnings rise accordingly. A rise (or fall) in hours of work causes a change in annual earnings. This may apply to approximately one-half of all paid workers. Other workers are paid on some other basis, such as an annual salary. How does one interpret an apparent association between changing hours of work and changing annual earnings for these workers. Changing hours of work do not necessarily cause a change in annual earnings among, say, managers or professionals, although over the long run there is an observed association in the changes between the two. It may be that the same economic forces that cause the relative annual earnings of these workers to rise also cause their relative hours worked to increase. If they move together, relative hourly wages would not change. Thus, for some workers it is not that changes in hours worked directly cause a change in annual earnings, but rather that forces are brought to bear on the two variables so that they move in the same direction.

Interpreting the Results in a Supply/Demand Framework, and Differences with the United States

These observations suggest that the labor market adjustments to economic forces that led to rising earnings inequality have been at least as much through a quantity (hours worked) as through the price (wages paid) adjustment; probably the quantity adjustment has dominated. This seems to be true for the early 1990s as well as the 1980s, although increases in earnings inequality do not appear to be as significant in the 1990s as the 1980s.

Forces on both the supply and demand side of the labor market can contribute to the change in earnings inequality. In their study of rising earnings inequality in a number of countries, Freeman and Katz (1994) conclude that "changes in the supply and demand for labour skills substantially alter wages and employment of different groups of workers in the manner predicted by economist's supply and demand market clear-ing model." That is, the supply–demand model works. They go on to note, however, that supply and demand factors by themselves cannot explain all of the differing changes in inequality among advanced countries. They note that demand factors are probably more or less the same in the developed countries, as they compete in the same global market with the same technologies. Supply factors may differ somewhat more, as demographics may be different and the development of the education and training systems may vary across countries.

They also argue, however, that something beyond supply and demand is needed to understand significant changes in labor markets. Knowledge of labor market institutions, which vary across countries, is also necessary. They note that "the stronger the role of institutions in wage determination, the smaller the effect of shifts in relative wages and, as a consequence, the greater will be their effect on relative employment." Earlier studies have noted that the evidence that is available suggests that in the United States, the rise in earnings inequality was associated more with changing relative wages than hours worked. In Canada, changing hours appear to dominate. This may be related to the larger role played by institutions in the Canadian labor market.

Unionization of the U.S labor force fell significantly during the period of interest here, while the change was much less significant in Canada (Card and Freeman 1994; Riddell 1993). Lemieux (1993) found that the difference in the unionization trends were associated with the lower rate of increase of labor market inequality in Canada than the United States. These same trends may also be associated with the tendency for changes in hours to play a more important role in Canada. Pressures for adjustment to changes in relative demand for labor may be reflected more through a redistribution of hours worked rather than hourly wages in a relatively stronger unionized setting.

But hours worked and wages paid respond to supply side influences as well. As noted in Freeman and Katz (1994), more generous income maintenance or unemployment insurance benefits may allow workers to be more rigid in the face of potential wage cuts, and reduce their willingness to take lower wages to obtain work, thus reducing supply-side pressures for wage cuts. This could contribute to the result that the adjustment is more on the hours than wage dimension.

In a related work, Kuhn and Robb (1996), when looking for an explanation of the rising unemployment gap between Canada and the

United States, noted that over the 1977–1992 period, the work behavior (in terms of weeks worked among prime-age males) changed very differently in the two countries. They observed that over this period work behavior changed such that Canadian men reduced their weeks of work, while Americans increased theirs. Some of this difference may be due to the fact that the 1990–1992 recession was much more severe in Canada than the United States, but it may also be reflective of differences in the changes in working time in the two countries.

In a simple supply–demand model, if labor demand for, say, lower paid and skilled workers decreases, shifting the demand curve to the left, one might expect to see an adjustment in both hourly wages paid and hours worked. The extent to which such a demand shift is reflected in one or the other would depend in part on the institutional arrangements in the country, as noted above. If there was a shift to the right in the demand for more highly skilled and paid workers, one would expect to see both their hours and wages rise. Data presented here suggest that both may have occurred, with the emphasis on changes in hours worked.

Other Factors Influencing Hours of Work

But aside from institutional arrangements, it may not be too surprising that the effects of decreases in demand for labor for, say, less-skilled and lower paid workers are reflected in changing relative hours of work. Economists have noted the reluctance of wages to adjust downward for some time. Hall (1995) talks about the unwillingness of employers to renegotiate wages of their workers to save jobs (or hours of work). He refers to work by Benley (1994) that documents the absence of renegotiation of wages in a depressed local labor market. By far the most common reason given by employers for this practice is that lowering wages would reduce morale, and hence presumably productivity. The notion that wages do not easily adjust downward, requiring a quantity response, is not new.

There are other incentives that employers may have for adjusting hours of work. For example, there is much talk of the desire for increased flexibility of employment levels in the face of changing product demand. Employers may use more part-time, contract or temporary workers, which would allow employment (and hours) levels to be more easily adjusted when the company faces a downturn, thereby reducing labor costs. This would influence weeks or hours worked, and earnings of many workers, increasing polarization of the hours variables. While it is not known to what extent this is actually occurring, it could be part of the explanation of these findings. Another often discussed incentive relates to the impact of fringe benefits and payroll taxes on hiring practices. At higher wage rates, the marginal cost to the firm in terms of UI or Q/CPP payroll taxes of longer hours of work is zero. However, engaging a new employee or extending the hours of lower paid employees does have a cost. Similarly, extending hours of work (and annual earnings) does not increase the cost of many fringe benefits, and this may be particularly important among the more highly paid. These possibilities may also encourage employers to adjust hours worked.

The Declining Relative Wages among the Young

While changing hours may play an important role in rising inequality in general, it does not do so for one particular dimension of the inequality story, the declining real and relative annual earnings among younger workers. This does not mean that there were not significant changes in hours worked among the young men. It means that changes in hours worked were very similar among younger and older workers, and hence relative values did not change. But relative wage rates did decline among younger workers.

Why would changing working time play a significant role in rising earnings inequality in general, but hourly wages dominate the increasing earnings gap between younger and older workers? We are discussing changes in relative earnings (or inequality) *among* groups, in this case by age. Changing relative earnings among age groups is one part, but only a small part, of the overall rise in earnings inequality. Increasing disparity *within* groups (defined in various ways) plays a larger role (Levy and Murnane 1992; Morissette, Myles, and Picot 1994; Richardson 1994).

These results indicate that there is real wage adjustment taking place in the Canadian economy. It may be observed more among young workers than elsewhere for a number of reasons. The first relates to relative education levels. Traditionally the young have enjoyed much higher levels of education than older workers. Recently this advantage has largely disappeared, as the educational attainment of older workers has risen quite rapidly. This is a result of the aging of the more highly educated "baby boom" generation. This decline in relative educational level among the young (relating to older workers) seems to account for around 30 percent of the earnings gap between the young and the old (Kapsalis, Morissette, and Picot 1997). Falling relative education levels would be associated with decline in wages.

A second reason relates to where labor market adjustment takes place. Adjustment of almost any type is typically more concentrated among younger workers. This includes migration, adjusting to changing regional economic circumstances, changes in skill acquisition in the face of changing demand, and wage adjustment. Among workers in general, the reluctance of wages to adjust downward was noted. This unwillingness on the part of both employees and employers to see wages adjust downward would be particularly strong among older employees, where an implicit contract with no expectations of downward wage adjustment may have developed over many years between the employee and the employer. Seniority provisions may also make such adjustment more difficult. And if downward wage adjustment, when it does occur, takes place in the open labor market (as opposed to internal labor markets in a company), this would have more impact on the young, as they are more frequently exposed to the pressures of the open labor market as they separate from firms more frequently.

It is likely easier for companies to adjust the wages of entry- (or near entry) level jobs downward in the face of decreased labor demand than it is to adjust the wages in jobs filled by experienced workers for reasons given above. This would influence the wages of predominantly younger workers.

For all these reasons, older workers are likely to be relatively immune to the downward adjustment of wages, at least relative to younger workers. Thus, when confronted with decreased labor demand for some particular group (say, less-skilled workers), companies may choose in general to adjust hours of work (and wages to a lesser extent) for a variety of reasons. These might include institutional factors, efficiency wage arguments, and the desire not to reduce moral and productivity, and the desire to have more flexibility in working time to keep labor costs low in the face of decreased product demand, and possibly incentives associated with the structure of payroll taxes. But some wage adjustment obviously takes place, and it appears to be concentrated among the young.

Notes

The author thanks Wendy Pyper for her usual excellent research assistance in preparing this paper.

1. Calculations were made with the self-employed who have employment earnings left in the sample, and the influence on the results were minimal.
2. Students could not be excluded from all of the data sets.
3. Since we do not have data for comparable points in the business cycle, we do not know if the excluded population is increasing in size. If it were, we would be underestimating the *change* in the distribution of hours worked. The increase in the share of labor force members at the bottom of the hours worked distribution would be underestimated, as would the impact of changes in hours worked on earnings inequality. Given the magnitude of the estimated population, however, we do not believe including this group would significantly change the findings.
4. In the upper deciles there can, of course, be little increase in weeks worked, as the average is very close to 52 weeks. Any increase in hours worked per year must come through increases in hours per week, not weeks per year. No decline in weeks worked was observed in the upper deciles.
5. Any difference in observed patterns between this period (1984–1993) and the earlier 1981–1989 period really reflects differences in the patterns of inequality in trends in the two recessions. This is because the expansionary period, 1984–1989, is common to both periods.
6. See the earlier section "The Data" for a listing of these surveys.
7. The major job is that with the most hours per week in the month of December. Data on weekly earnings, hours per week, and hourly wages were also computed using all jobs held during the year, weighted according to the number of hours worked in each job (for all years except 1984). A comparison was made between the distributions and the change in the distributions observed using only the December job and all jobs held during the year, and the trends were the same for both approaches.
8. This is usual hours worked per week on the LFS, which includes unpaid overtime, and overtime worked on a regular basis. Similar data are obtained from the special surveys used here, including the SWH, SUM, LMAS, and SLID. In these surveys, usual hours worked as also collected (in SLID and LMAS the reference is paid usual hours). Data are collected for usual hours per day, and days per week, except for the 1993 SLID survey, which collects data on the usual hours per week directly. These surveys also show a decline in the proportion of males

working regular (35–40) hours, from 75 percent in 1981, 72 percent (84), 71 percent (86), 67 percent (89), and 66 percent (93).

9. This increase in the share of men working longer hours was not observed in the LMAS/SLID data. This is likely due to the differences in the manner in which people reported hours worked in the two surveys. The LFS is a consistent source and likely more reliable.
10. When a respondent is asked, "What was your usual wage of salary before taxes and other deductions from this employer," they can respond in terms of wages per hour, week, month, or year. In the special surveys, from 36 percent (SWH) to 53 percent (SUM) reported hourly wage directly. For the remainder, hourly wage is computed by dividing earnings over the period (for example per week or per year) by the usual hours worked.
11. Among all males, the result is not quite as striking. Inequality in weekly earnings rose 14 percent between 1981 and 1989, according to these data, and 37 percent of this increase was associated with the change in hours per week, an additional 53 percent by the covariance term, which implies that the tendency of higher paid workers to work longer hours increased over the period.
12. It should be noted, however, that the rise in the inequality of weekly earnings observed in these data sets for this 1984–1993 period is not reflected in the SCF data, where weekly earnings inequality does not increase significantly over the 1984–1993 period. In the SCF data, inequality in weekly earnings, after increasing 25 percent among males in the 1981–1988 period, is seen to rise little over the 1984–1993 period. This leveling is not observed in the special surveys data. Weekly earning inequality rises around 15 percent during the 1980s (comparable to the SCF results), but 15–20 percent over the 1984–1993 period, which is much higher than in SCF, where virtually no increase was observed. As a time-series measure of weekly earnings, the SCF has a major advantage, it is a consistent series of measures from the same survey vehicle. The special surveys, which are similar in the way in which they treat hours and wages, do differ in some ways. This is particularly true for the 1993 observation from SLID. Hours per week are measured in a slightly different manner (SLID measures usual hours per week directly, while the previous survey measured usual hours per day and days per week). It seems likely that the consistent series from the SCF would be a more reliable source of trends in weekly earnings than that observed in the special surveys.
13. Trends in the change in average weekly earnings were somewhat different from the two sources. The decline in average weekly earnings at the bottom end of the distribution observed for the 1981–1989 period in SCF (Tables 2 and 5) is evident in the data from the special surveys for all males, but not for prime-age males. The change in the *distribution* of weekly earnings is similar to the two sources (both show rising inequality), but the change in the *level* is different. This is probably because the level of weekly earnings in the 1981 SWH appear to be low relative to the other special surveys. This would cause the growth to be overestimated between 1981 and 1989. Average weekly earnings from the special surveys were

compared with those from SCH. For all males, for example, the ratio of weekly earnings in the special surveys compared to those in SCF was in the 97 percent to 101 percent range for all years (84, 86, 89, 93) except 1981, when it was 93 percent. A similar pattern was observed for prime-age males. Average weekly earnings appear to be underestimated in the 1981 SWH.

14. For example, among men 25–44, the employment/population ratio fell from 90.4 percent in 1981 to 88.2 percent in 1989, and 83.3 percent in 1995. Among women of the same age, the trend was up through the 1980s from 60.8 percent in 1981, to 70.8 percent in 1989, to 70.6 percent in 1995.
15. The special surveys data also show a decline in the share of prime-age women working a regular 35–40 hour week, from 65 percent in 1981, to 63 percent in 1989, and 60 percent in 1993.
16. That is, those with lower wages are increasingly working shorter hours, and those with higher wages are increasingly working longer hours.
17. Earlier work shows that among full-time, full-year, female workers, earnings inequality rose in the 1980s, and declined somewhat in the 1990s (Morissette, Myles, and Picot 1994; OECD 1996).
18. Earlier work for the 1980s showed that among women working full-time, full-year earnings inequality rose, but when part-year and part-time working women are included, it fell (Morissette, Miles, and Picot 1994).
19. The one exception is among the 17- to 24-year-olds during the 1984–1993 period, where relative hours worked per week (relative to 45- to 54-year-olds) decline 12 percent compared to a decline in hourly wages of 18 percent.

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5

Patterns of Foregone Potential Earnings among Working-Age Males, 1975–1992

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Fundamental changes in labor market patterns among U.S. prime-age men over the past two decades have been the focus of numerous recent research studies and media accounts. Increases in wage inequality and in male joblessness are the most important of these changes; assertions of an increase in part-time and "contingent" work have also been made. In addition, there is evidence of a more general decline in the total annual hours of market work of the typical working-age male.[1]

In this chapter, we use a new statistical indicator, *foregone potential earnings* (FPE), to measure the extent to which the prime-age male population (civilian nonstudent 18- to 64-year-old males) underutilizes its human capital. We define the annual value of an individual's human capital[2] to be the amount that an individual would earn if he worked full time, full year (FTFY); that is, 52 weeks per year and 40 hours per week. This amount is the individual's *potential earnings*. FPE is the gap between an individual's actual earnings and his potential earnings and is thus an indicator of the underutilization of human capital.[3]

We use our FPE indicator to examine trends in human capital underutilization for the entire population of working-age males, and for various population subgroups, during the 1975–1992 period. We also examine trends in the reasons given for the failure to fully utilize human capital. We find that over the 1975–1992 period, per capita FPE increased by almost 3 percent for all working-age males. This increase stems from a 12 percent decrease in real per capita earnings and a 10 percent decrease in the real per capita potential (or FTFY) earnings of these individuals. When we aggregate the reasons given for underutilization into exogenous constraints on working and individual preferences for not working, we find that the share of FPE attributable to the former has declined, while FPE attributable to the latter has risen. This shift is particularly pronounced for older, less educated, nonwhite men.

Our FPE indicator provides a more complete picture of economic performance than other statistical measures of labor force activity, such as the unemployment rate. Whereas the unemployment rate simply indicates the percent of individuals in the labor force looking for work, our FPE indicator applies to individuals in and out of the labor force, quantifies in dollar amounts the level of underutilization, and identifies the sources of underutilization. For example, our FPE indicator demonstrates the increased importance of retirement relative to unemployment as a source of underutilization. Similarly, it can be used to measure the effect of policy changes. For example, what happens to underutilization due to illness and disability when federal health care policy changes? It can also be used as a supplemental indicator for assessing the macroeconomic performance of the economy, measuring both the extent and composition of slack in the utilization of the nation's labor resources. In essence, FPE provides the first measure that values in monetary terms the level and trend of human capital underutilization.

TRENDS IN HOURS WORKED, 1975–1992

Figure 1 shows the trend in average annual work hours for the white, nonwhite,[4] and total male working-age population over the 1975–1992 period, as reflected in the March supplement to the annual Current Population Survey (CPS).[5] For both racial groups, mean annual hours decreased during the 1980–1983 recession; the subsequent recovery failed to return this value to its pre-1980s level for either racial group. Over the entire period, the trend of annual work hours is slightly negative for all working-age males and for the two racial subgroups.

Figure 1 **Mean Annual Hours Worked, 18- to 64-Year-Old Males, 1975–1992**

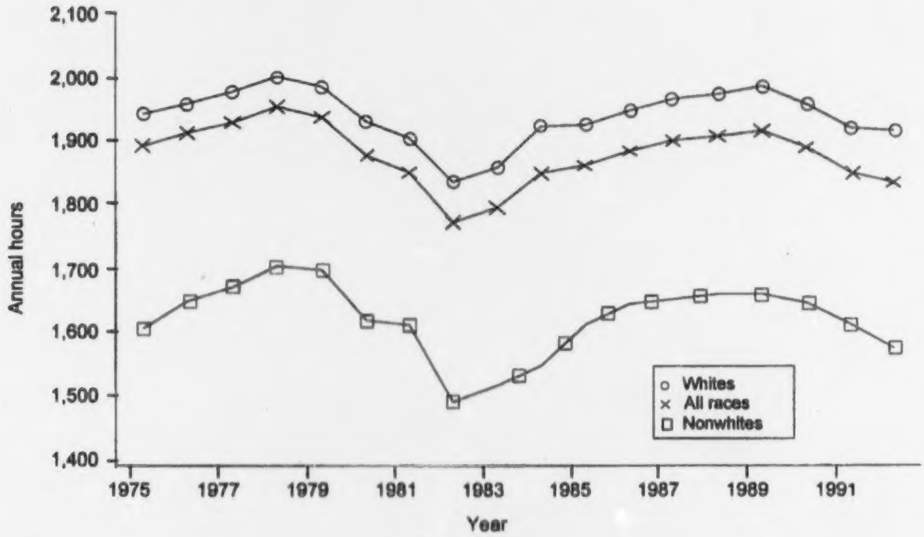


Table 1 indicates the reason for the decrease in this average value. It shows the percentage of the sample in four annual hours-worked categories for the paired recession years of 1975 and 1991 and the paired cyclical peak years of 1979 and 1989. The most noteworthy change is the 26 percent increase in the proportion of jobless males (those with zero work hours) over the 1975–991 period—an increase from 7.7 to 9.7 percent of the working-age population. However, the share of working males employed less than 2,080 hours per year decreased by about 6 percent for the paired recession years and 12 percent for the paired peak years. The share of prime-age males working exactly at capacity declined 5 percent over the recession years and remained constant over the peaks, while the share working in excess of capacity increased by about 6 percent for both pairs of years. Over the sets of paired years, then, there has been a hollowing out of the middle of the annual hours distribution, with an increase in the mass at both extremes.[6] These hours-worked trends suggest substantial shifts in labor supply and demand over the period. Although the pattern of changes in the mean and variance in male earnings have been extensively studied, including changes in the level and distribution of both

Table 1 Percentage of 18- to 64-Year-Old Males in Annual Hours Worked

Year	0 hr.	1–2,079 hr.	2,080 hr.	>2,080 hr.
Recession years				
1975	7.7	31.1	34.6	26.6
1991	9.7	29.2	32.8	28.3
Change (1991–1975)	2.0	–1.9	–1.8	1.7
Peak years				
1979	7.4	29.7	34.2	28.7
1989	8.8	26.1	34.7	30.4
Change(1989–1979)	1.4	–3.6	0.5	1.7

SOURCE: Authors' calculations, March 1976, 1980, 1990, 1992 CPS.

wage rates and hours worked, the sources of the observed shifts remain little understood.[7]

HUMAN CAPITAL UNDERUTILIZATION OF WORKING-AGE U.S. MALES

The Concept and Estimation of FPE

We define the earnings associated with full use of human capital as potential earnings (PE) and measure this value as the product of an individual's predicted wage rate[8] and 2,080 hours. The individual's earnings are measured as the product of the actual number of hours that the person works in a year and his predicted wage rate. FPE, then, is the number of dollars that an individual's earnings fall short of PE.[9] It can be thought of as weighted foregone hours—hours worked less than the norm—where the weight is based on an estimate of the value of the person's productive capabilities in the labor market. For any group of working-age males, *I*, we measure FPE as an average value,

$$FPE_I = \frac{\sum_{i \in I} Potential\ Earnings_i}{N} - \frac{\sum_{i \in I} Earnings_i}{N},$$

where N is the number of individuals in I. So defined, FPE measures the extent to which the utilization of human capital deviates from a socially accepted norm of full-capacity utilization; in this case, 2,080 hours per year.[10]

Our measure of FPE neglects the role of important nonmarket activities in two ways. First, we assume that human capital is utilized only through paid market work. While human capital is also utilized through nonmarket activities, our purpose is to analyze trends in potential and foregone potential earnings. Since nonmarket activities are, by definition, unpaid, they cannot contribute to the realized earnings of an individual; hence, we neglect them here.

Second, we ignore the fact that certain nonmarket activities, such as child care, must be performed. If the individual does not perform them, he or she must obtain services from someone else, perhaps by purchasing them. Thus, FPE may not represent the *net* increase in either individual realized earnings or in aggregate earned income when the individual moves to full utilization.[11]

FPE of Working-Age Males

We begin our examination of FPE with Table 2, which shows the trends in various earnings measures for the working-age U.S. male population.[12] Over the 1975–1992 period, aggregate real earnings increased from $1.26 trillion to $1.47 trillion, or 17 percent.[13] During this same period, the total male working-age population grew from about 52 million to about 69 million, or 32 percent. Hence, per capita earnings for working-age males fell nearly 12 percent over the period, from about $24,000 to $21,000, which is consistent with other estimates of sagging mean earnings of male workers.

We estimate that over the same period, aggregate potential earnings of all working-age males in the United States rose from $1.48 trillion to $1.77 trillion, an increase of 19 percent. However, because of the 32 percent growth in the size of the working-age male population over this period, per capita potential earnings fell from $28,206 to $25,494, a decrease of 9.6 percent.

By comparing the level of per capita earnings to per capita potential earnings, we can measure the extent to which working-age males fail to utilize their stock of human capital. Over the 1975–1992 period, the gap between aggregate earnings and aggregate potential earnings (aggregate FPE) increased from \$.22 trillion to \$.30 trillion, or 36 percent. In per capita terms, FPE increased nearly 3 percent, from \$4,201 to \$4,313.[14]

THE REASONS FOR FOREGONE POTENTIAL EARNINGS

Self-Reported Reasons for FPE

Table 2 shows that per capita real FPE ranged from about \$3,800 in 1978 to over \$5,000 in the recession year of 1982. From respondents'

Table 2 Per Capita Earnings Measures, 18- to 64-Year-Old Males[a]

Year	Earnings (\$)	Potential earnings (\$)	Foregone potential earnings (\$)
1975	24,004	28,206	4,201
1976	24,630	28,780	4,150
1977	24,367	28,261	3,893
1978	24,966	28,801	3,836
1979	24,849	28,634	3,785
1980	24,039	28,725	4,236
1981	22,996	27,335	4,339
1982	22,380	27,424	5,045
1983	22,303	27,295	4,992
1984	22,919	27,448	4,529
1985	23,011	27,310	4,299
1986	23,892	28,329	4,437
1987	23,793	28,101	4,308
1988	23,373	27,317	3,944
1989	23,333	27,153	3,820
1990	22,285	26,176	3,891
1991	21,450	25,613	4,163
1992	21,181	25,494	4,313
Change 1975–1992 (%)	−11.8	−9.6	+2.7

SOURCE: Authors' calculations, March 1976–1993 CPS.

[a] All dollar amounts are adjusted to 1993 dollars using the CPI-U-X1 cost index.

answers to questions regarding why they work less than the FTFY norm, per capita FPE for each year can be decomposed into the following comprehensive set of "reasons":[15]

- work is not available (unemployed),
- discouraged from seeking work,
- illness/disability,
- retirement,
- voluntary part-time work,
- housework, including child care, or
- other.

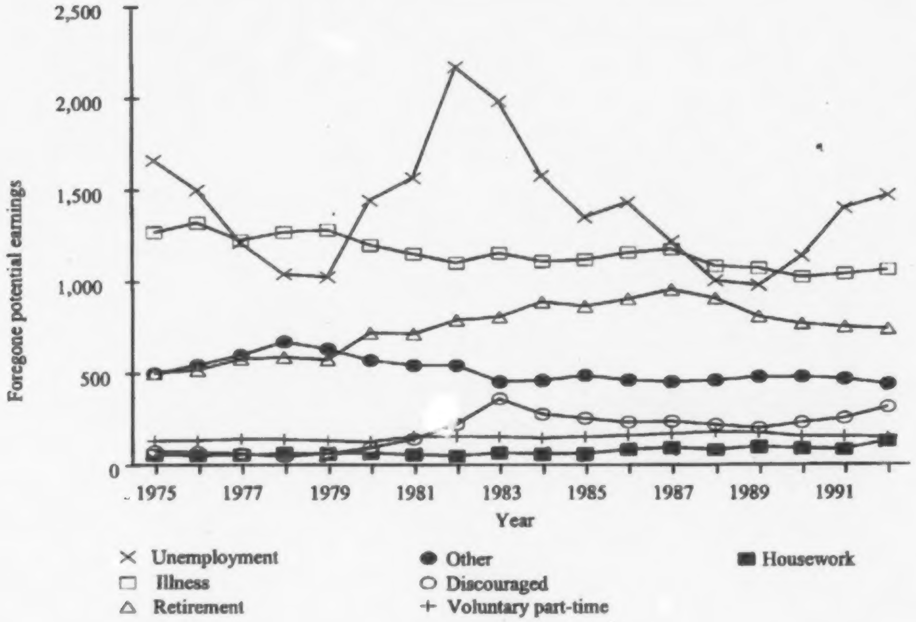
The level and trend of these components of per capita FPE are presented in Figure 2 for the 1975–1992 period. The vertical sum of the component values for each year equals per capita FPE.

With the exception of the late 1970s and late 1980s booms, a lack of employment opportunities for those seeking work is the largest component of FPE. This unemployment component peaked during the recession of the early 1980s, when it accounted for nearly $2,200 per person of FPE, and was at its lowest at the end of the expansion of the late 1980s, when it fell to less than $1,000 per person. Over the period, per capita FPE due to unemployment shows a slight downward trend of about $120 per decade.[16]

The second component of FPE is labeled "discouraged workers," and it too reflects macroeconomic conditions. The value of this discouraged worker effect ranged from a low of about $100 per person (or about 2 percent of total FPE) at the end of the 1970s, to a high of nearly $400 (nearly 6 percent of the total) during the early-1980s recession. While this value declined during the expansion of the 1980s, it never fell below $200 per person, and rose to over $300 by the end of the period. Due to this discouraged worker effect, per capita FPE showed an upward trend over the period of about $140 per decade.

Illness or disabling health conditions forms the second most important reason for FPE, and accounted for a per capita value of about $1,000 to $1,300 per year over the period. The trend in FPE due to this factor is clearly downward, however, at about $150 per person per decade. This downward trend in foregone earnings due to illness/dis-

Figure 2 The Per Capita Gap between Earnings and Potential Earnings, 18- to 64-Year Old Males, by Reason



ability contradicts a growing incidence of illness/disability problems among the working-age population reported in other studies.[17]

Retirement is the third most important reason for FPE, and accounted for $500 per capita to nearly $1,000 per capita. This source of FPE is the most rapidly growing among the set of reasons given by working-age males for the failure to fully use human capital. Per capita FPE due to retirement has grown about $190 per decade, or nearly $350 over the 1975–1992 period.

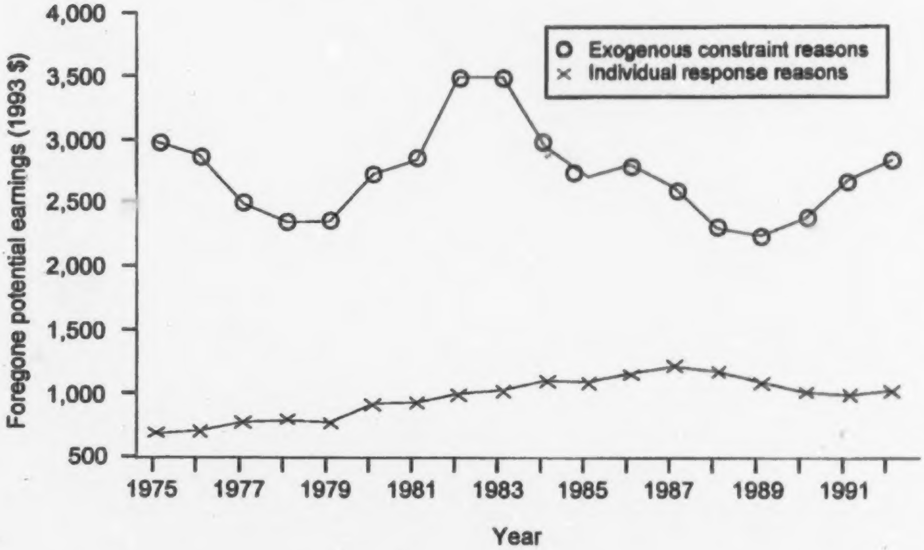
The remaining reasons for FPE (housework, voluntary part-time work, and other) account for a relatively small share of total FPE per person—ranging from 14–23 percent of the total over the period. Aggregate FPE attributable to this set of reasons has crept up slowly over the period.

Underutilization Due to Exogenous Constraints and Individual Response

The underutilization of human capital due to exogenous constraints placed on individuals carries quite different social and policy implications than that due to voluntary, individual choices. For this reason, we have divided per capita FPE into two components—that arising from individual responses to incentives (retirement, voluntary part-time work, and housework), and that stemming from exogenous constraints on the utilization of human capital (work not available, discouraged from seeking work, and illness).[18]

Figure 3 shows the level of per capita FPE due to exogenous constraint and individual response reasons for the working-age male population. An upward trend for individual response reasons is observed, while the trend for exogenous constraint reasons is negative. At the beginning of the period, FPE due to individual response reasons was 23 percent as large as exogenous constraint reasons for FPE; by the end of the period, the individual response reasons had grown to over 36 per-

Figure 3 Exogenous Constraint and Individual Response Reasons for Foregone Potential Earnings, 18- to 64-Year-Old Males



cent of the value of the exogenous constraint reasons. Over the 1975–1992 period, per capita FPE attributed to individual response reasons increased by about $240 per decade, while per capita FPE due to exogenous constraint reasons fell by about $130 per decade.

FOREGONE POTENTIAL EARNINGS PATTERNS AMONG RACE, AGE, AND EDUCATION SUBGROUPS

The overall patterns of working-age male human capital underutilization described above conceal substantial differences among race/age/education subgroups. In this section, we summarize a few of the more prominent of these differences.[19] We begin with a discussion of racial differences and then present differences among age and education subgroups.

Racial Differences in FPE

Over the 1975–1992 period, the ratio of nonwhite to white potential earnings fell from 0.74 to 0.71. The earnings potential of the mean white male fell by an average of $1,104 per decade; that for the mean nonwhite male fell by $1,188. As a result, the racial gap in potential earnings increased slightly over the period.[20] Table 3 shows 1975 levels of PE and FPE for both nonwhites and whites, along with the reasons for FPE in that year, and the constraint/response breakdown in the causes for FPE. It also summarizes trends in all of these categories expressed in "dollars of average per decade change" over the 1975–1992 period.

The most striking pattern is the decline in per capita FPE for nonwhites over the period, in contrast to virtually no change in per capita white FPE. The difference in the "Per decade change" columns implies that the racial gap in the FPE indicator of labor underutilization narrowed by nearly $240 over the 1975–1992 period—or by more than one-fifth of its initial level of about $1,170.

While the contribution of unemployment to underutilization or FPE decreased over the period for both racial groups, the decrease in the unemployment reason for FPE was larger for nonwhites than for

Table 3 Foregone Per Capita Potential Earnings, 18- to 64-Year-Old Males, by Race[a] ($)

	Nonwhites		Whites	
	1975 Level	Per decade change	1975 Level	Per decade change
Potential earnings	21,663	–1,188	29,400	–1,104
Total FPE	5,190	–129	4,021	3
Unemployment	2,163	–264	1,569	–118
Discouraged	174	240	59	102
Illness	1,838	–196	1,167	–163
Housework	74	46	55	27
Retirement	230	130	551	229
Voluntary P/T	120	14	138	24
Other	592	–103	481	–98
Individual response FPE	424	190	744	280
Exogenous constraint FPE	4,174	–216	2,795	–180

SOURCE: Authors' calculations, March 1976–1993 CPS.

[a] All dollar amounts are adjusted to 1993 dollars using the CPI-U-X1 cost index.

whites. The failure of the economy to perform at full capacity appears to have taken a smaller toll on nonwhites (relative to whites) at the end of the period than it did at the beginning.

FPE due to being discouraged from seeking work is quantitatively small relative to FPE due to unemployment. However, this discouraged-worker FPE was three times larger for nonwhites than for whites at the beginning of the period, and increased at twice the rate for nonwhites relative to whites over the period.

For both racial groups, a large share of the decrease in capacity utilization is attributable to the increase in pre–age-65 retirement. Per capita "early" retirement FPE for whites was double that for nonwhites at the beginning of the period and increased more rapidly over the period. Primarily because of the more rapid growth in FPE due to retirement for whites, the individual response reasons for FPE grew more for whites over the period than for nonwhites.

Age Differences in FPE

Table 4 shows the trends in potential earnings and FPE for the youngest (ages 18–24) and oldest (ages 55–64) groups, as most of the interesting patterns are concentrated in these groups. Per decade, the earnings potential of 18- to 24-year-olds fell by $2,700, while mean potential earnings of older working-age males decreased only $960. Over the entire period, the ratio of the potential earnings of the youngest group to that of the oldest group fell from 0.64 to 0.52—a radical drop of 12 basis points.

FPE is higher for the older group than for the younger group, which is not surprising given the substantially higher potential earnings of the older group. Moreover, the old-to-young gap in FPE has been rising over time. Over the entire 1975–1992 period, per capita FPE for

Table 4 Foregone Per Capita Potential Earnings for the Youngest and Oldest Age Groups[a] ($)

	Age 18–24		Age 55–64	
	1975 Level	Per decade change	1975 Level	Per decade change
Potential earnings	17,645	–2,700	27,725	–960
Total FPE	4,207	–379	7,369	1,130
Unemployment	2,592	–503	1,133	–29
Discouraged	218	193	62	126
Illness	283	7	2,985	–602
Housework	27	39	71	11
Retirement	1	7	2434	1,562
Voluntary P/T	357	19	235	129
Other	729	–142	448	–67
Individual response FPE	385	66	2,740	1,702
Exogenous constraint FPE	3,093	–303	4,181	–506

SOURCE: Authors' calculations, March 1976–1993 CPS.

[a] All dollar amounts are adjusted to 1993 dollars using the CPI-U-X1 cost index.

youths fell by almost $700, while per capita FPE for the group of older workers rose by over $2,000.

For youths, the $700 drop in FPE has been driven by a decrease of more than $500 per decade in FPE attributable to reduced unemployment-generated nonutilization. However, the sizable increase in FPE among youths due to the discouraged worker effect—about $350 over the period, a twofold increase—is disturbing and runs in the opposite direction to the unemployment effect. The very large increase in FPE for the older age group—over $2,000 during the 18-year period—is more than explained by the rapid increase in retirement over the period. However, the retirement-induced increase in FPE for this older group was offset by a substantial decrease in the amount of foregone earnings due to illness/disability; from an average of about $3,000 per year in FPE at the beginning of the period, to about $2,200 by the end of the period.[21]

Education Differences in FPE

Table 5 summarizes the pattern of potential earnings and the utilization of this potential over the 1975–1992 period for the two lowest education groups—dropouts and those with a high school degree but no college. Over the 18-year period, potential earnings for both groups fell dramatically. The average per decade decrease in potential earnings is $4,265 for dropouts and $3,571 for high school graduates. Of the four education groups, only college graduates showed an increase in potential earnings over the period (not shown in table). The increasing return to years of schooling is clearly seen in the widening gap in potential earnings among the education groups, even between high school dropouts and those with a terminal high school degree.

Per capita FPE for the high school dropouts decreased slightly over the period, by about $80 per decade, while FPE for those with a high school degree increased about $200 per decade. The reasons for the level of and change in FPE for these low-education groups are dominated by unemployment, discouragement over finding work, and illness. Earnings foregone due to unemployment decreased for both low-education groups over the period. However, per capita FPE due to the discouraged worker effect increased by $264 per decade for the group of dropouts and by nearly $100 per decade for those with a high school

Table 5 Foregone Per Capital for Those with No College[a] ($)

	High school dropouts		High school graduates	
	1975 Level	Per decade change	1975 Level	Per decade change
Potential earnings	22,280	–4,265	27,491	–3,571
Total FPE	5,901	–81	3,865	202
Unemployment	2,023	–137	1,787	–132
Discouraged	103	265	84	154
Illness	2,548	–302	905	85
Housework	73	40	46	37
Retirement	550	138	481	147
Voluntary P/T	104	7	110	5
Other	500	–91	452	–94
Individual response FPE	727	185	636	189
ExogenousconstraintFPE	4,674	–175	2,777	107

SOURCE: Authors' calculations, March 1976–1993 CPS.

[a] All dollar amounts are adjusted to 1993 dollars using the CPI-U-X1 cost index.

degree. A large increase in underutilization due to retirement is also recorded for both groups.

FOREGONE POTENTIAL EARNINGS PATTERNS FOR VULNERABLE GROUPS

The patterns discussed in the previous section reveal substantial variation in human capital underutilization among subgroups of the male working-age population. In general, nonwhite youths and older males—especially those with low schooling levels—have the highest levels of underutilization. These same groups display the largest increases in human capital underutilization over time. Here, we focus on the youngest and oldest nonwhite groups with the lowest schooling levels, and compare their FPE patterns with those of the average male in their age group and the average working-age male, irrespective of age.

Low-Education Minority Youths

Consider, first, low-education minority youths (Table 6). The top row of the table, potential earnings, shows vividly the declining prospects of low-education minority youth. Over the 18-year period, real potential earnings fell by 16 percent per decade for both nonwhite youths who dropped out of high school and those with a terminal high school degree. This compares with a 15 percent decadal drop for all youths and a 5 percent drop for all males.

For both low-education groups of minority youths, FPE fell over the period. However, the decrease in FPE must be interpreted in the context of a decreasing level of potential earnings. The ratio of per capita earnings to total potential earnings, which reflects the percent of potential earnings utilized, fell over the period for both low-education minority groups, by over 3 percentage points for the dropouts and 4 percentage points for the terminal high school graduates.

For all of the groups, unemployment accounted for the largest portion of unused earnings potential. Although this share fell over the period for all four groups, discouragement over finding work accounted for an increasing share of FPE. For all the groups, the individual response reasons for FPE increased over the period.

Older, Low-Education Minority Males

Table 7 shows that potential earnings decreased substantially for older, low-education minority workers (by 9 percent for dropouts and 6 percent for high school graduates, per decade), relative to both all older working-age men (3 percent) and all males (5 percent).

Similarly, our indicator of the underutilization of human capital—FPE—is very high for older, low-education minority males, especially relative to their earnings potential. At the beginning of the period, these groups utilized only about 60–65 percent of their earnings potential, compared to 73 percent and 85 percent for all older workers and all males, respectively. However, compared to low-education minority youths, FPE for the older, low-education minority workers rose substantially over the 1975–1992 period.

The reasons for FPE among nonwhite, low-education, older males are dominated by unemployment, retirement, and illness. For both

Table 6 Foregone Per Capita Potential Earnings, 18- to 64-Year Old Males[a] ($)

	Nonwhite dropouts, ages 18–24		Nonwhite high school degree, ages 18–24		All males, ages 18–24		All working-age males	
	1975 Level	Per decade change	1975 Level	Per decade change	1975 Level	Per decade change	1975 Level	Per decade change
Potential earnings	14,210	−2,221	16,475	−2,607	17,645	−2,700	28,206	−1,518
Total FPE	6,134	−721	4,846	−455	4,207	−379	4,201	17
Unemployment	3,085	−877	3,090	−717	2,592	−503	1,661	−122
Discouraged	719	324	323	332	218	193	76	140
Illness	793	−91	365	52	283	7	1,271	−150
Housework	90	77	63	38	27	39	58	32
Retirement	0	15	0	1	1	7	502	189
Voluntary P/T	212	10	316	1	357	19	135	19
Other	1,236	−179	688	−162	729	−142	499	−91
Individual response FPE	302	102	379	40	385	66	695	240
Exogenous constraint FPE	4,597	−643	3,778	−333	3,093	−303	3,008	−132

SOURCE: Authors' calculations, March 1976–1993 CPS.

[a] All dollar amounts are adjusted to 1993 dollars using the CPI-U-X1 cost index.

Table 7 Foregone Per Capita Potential Earnings, 18- to 64-Year-Old Males[a] ($)

	Nonwhite dropouts, ages 55–64		Nonwhite high school degree, ages 55–64		All ages, 55–64		All working-age males	
	1975 Level	Per decade change	1975 Level	Per decade change	1975 Level	Per decade change	1975 Level	Per decade change
Potential earnings	17,607	–1,610	24,977	–1,626	27,725	–960	28,206	–1,518
Total FPE	7,201	134	8,019	1,438	7,369	1,130	4,201	17
Unemployment	1,261	–131	1,673	–143	1,133	–29	1,661	–122
Discouraged	30	228	0	248	62	126	76	140
Illness	4,616	–598	3,158	96	2,985	–602	1,271	–150
Housework	111	–4	110	43	71	11	58	32
Retirement	770	686	2,111	1,387	2,434	1,562	502	189
Voluntary PT	161	27	255	121	235	129	135	19
Other	252	–74	712	–314	448	–67	499	–91
Individual response FPE	1,042	709	2,476	1,551	2,740	1,702	695	240
Exogenous constraint FPE	5,906	–501	4,831	201	4,181	–506	3,008	–132

SOURCE: Authors' calculations, March 1976–1993 CPS.

[a] All dollar amounts are adjusted to 1993 dollars using the CPI-U-X1 cost index.

groups—and for all older males—illness is the single largest reason for FPE; in 1975, it accounted for nearly two-thirds of FPE for the dropout group, and 40 percent of FPE for the older workers with a terminal high school degree.[22] For the dropout group, FPE attributed to illness declined over the period, as it did for the two comparison groups. For all of the older groups, retirement accounted for an increasingly large share of FPE over the period, while unemployment as a reason for FPE declined. It is noteworthy that nonwork due to the discouraged-worker effect accounted for very little of FPE for the nonwhite, low-schooling older group at the beginning of the period; however, this source of FPE grew rapidly over the period for this vulnerable population.

Largely because of the increase in retirement, individual response reasons for FPE grew for all of the older groups. This growth, in combination with the decrease in potential earnings, caused the percentage of potential earnings realized by older, low-education minority males (not shown) to fall substantially over the period, by 14–16 percentage points for the two low-schooling groups, compared with decreases of 9 percentage points for all older males and 1.6 percentage points for all males.

EXOGENOUS CONSTRAINT AND INDIVIDUAL RESPONSE REASONS FOR FPE

The patterns of underutilization described here raise the question of the extent to which the reduction in human capital utilization has derived from changes in the exogenous constraints that people face or in their individual responses to incentives. As we noted above, underutilization of human capital due to exogenous constraints placed on individuals carries quite different social and policy implications than that due to voluntary individual choices.

Over the 18-year period, an upward trend in individual response reasons for underutilization is observed, while the contribution of exogenous constraints to underutilization appear to be decreasing. At the beginning of the period, individual response reasons accounting for FPE were about 23 percent as large as those associated with exogenous constraints. However, by the end of the period, individual response reasons were 36 percent as large as the exogenous constraint reasons. Over the 1975–1992 period, per capita individual response reasons for underutilized human capital increased by about $240 per decade, while per capita exogenous constraint reasons fell by about $130 per decade.

Our calculations allow an even deeper assessment of these response/constraint sources of human capital underutilization among various age/race/education subgroups. In Table 8, we break the gap between earnings and potential earnings into the two components of individual response and exogenous constraint reasons, and show the ratio of these two values for the subgroups, for 1975 and 1992. We also show the percentage change in this measure over the two years for each of the subgroups. Overall, and for each of the subgroups, the individual response/exogenous constraint ratio increased rapidly over the 1975–1992 period. For all working-age males, the ratio rose by 57 percent. For the oldest individuals, the ratio increased by 122 percent, indicating the increasing importance of individual retirement decisions in explaining the growth in foregone potential earnings. Large per decade increases in this ratio are also recorded for older, nonwhite high school dropouts and graduates, and for young, nonwhite high school dropouts.

SUMMARY AND CONCLUSIONS

In this chapter we have defined a new indicator of the level of human capital, potential earnings, and a new indicator of labor underutilization, foregone potential earnings. FPE is the gap between the norm of full time-full year work and the hours a person actually works, weighted by his predicted hourly wage. We measure this value in 1993 dollars and interpret it as the amount of potential earnings that the individual foregoes. We have used this concept to assess the levels and trends of human capital and its utilization among U.S. working-age males from 1975 to 1992. Overall, the time-related patterns in both potential earnings and the utilization of this potential indicate that underutilization of the stock of male human capital has been increasing over the period. This trend in human capital underutilization has been concentrated among very young and old workers, those with the lowest

Table 8 The Levels and Percent Changes in the Ratio of Individual Responses to Exogenous Constraint Sources of Foregone Per Capita Potential Earnings[a]

	1975 Level of FPE ($)			1992 Level of FPE ($)			
	Individual responses	Exogenous constraints	1975 IR/EC ratio	Individual responses	Exogenous constraints	1992 IR/EC ratio	Change in IR/EC (%) 1975 to 1992
All working-age males	695	3,008	0.23	1,025	2,846	0.36	57
All nonwhites	424	4,174	0.10	757	3,748	0.20	102
All whites	744	2,795	0.27	1,106	2,573	0.43	59
Ages 18–24	385	3,093	0.12	495	2,563	0.19	61
Ages 25–39	140	2,318	0.06	227	2,437	0.09	55
Ages 40–54	340	3,126	0.11	476	3,227	0.15	34
Ages 55–64	2,740	4,181	0.66	4,961	3,387	1.46	122
High school dropouts	727	4,674	0.16	981	4,216	0.23	45
High school graduates	636	2,777	0.23	937	3,048	0.31	37
Some college	678	2,287	0.30	949	2,637	0.36	20
College graduate	778	1,321	0.59	1,263	1,822	0.39	17
Nonwhite dropouts, ages 18–24	302	4,597	0.07	405	3,435	0.12	68
Nonwhite high school graduates, ages 18–24	379	3,778	0.10	388	3,317	0.12	17
Nonwhite dropouts, ages 55–65	1,042	5,906	0.18	2,674	5,008	0.53	197
Nonwhite high school graduates, ages 55–64	2,476	4,831	0.51	4,798	4,843	0.99	94

SOURCE: Authors' calculations, March 1976–1993 CPS.

[a] All dollar amounts are adjusted to 1993 dollars using the CPI-U-X1 cost index.

education levels, and nonwhites. Finally, we note that while exogenous constraints on human capital utilization outweigh individual choices to underutilize, the relative contribution of choice-based FPE has increased over the period.

Notes

1. A December 1, 1994, front page *New York Times* story inquired, "So why are so many men—healthy men in the prime of life—working less than ever before?" (Nasar 1994). See also Buron and Haveman (1995), Buron, Haveman, and O'Donnell (1995), Freeman (1994), Katz and Murphy (1992), and Juhn (1992).
2. The human capital embodied in an individual is taken to be the value of the "bundle" of his characteristics—for example, schooling, skills, age, race, and health status—when fully used in productive economic activities. The independent effect of any one of these characteristics on the individual's observed (or estimated) wage rate is taken as an estimate of the market valuation of the hourly rental value of the characteristic. Hence, the market-determined "use-value" of an hour of the individual's work time—his wage rate—measures the economic value of an hour's worth of his human capital. This convention implies that the returns to race and gender found in human capital studies reflect real productivity differences and not discriminatory treatment of these traits in the labor market.
3. We assume those working full time, full year or more are using their human capital at capacity; no credit is given for work in excess of, 2,080 hours per year. While work patterns above 2,080 hours per year are also of interest, this chapter concentrates on underutilization of human capital, and hence, those individuals who work less than the full-time, full-year norm. Therefore, we cap each individual's work hours at 2,080 and count those with 2,080 or more hours of work as having zero unutilized hours.
4. "Whites" refers to white non-Hispanics; "nonwhites" are all others.
5. The standard method of calculating annual hours from the CPS is to multiply weeks worked in the last year by hours usually worked in a week. If reports of the latter correspond to modal hours rather than mean hours, as seems likely, this estimate is incorrect. In this analysis, we adopt a different convention and employ information on weeks worked part time and hours worked last week in the estimation of annual hours for some individuals. If an individual usually works full time (i.e., at least 35 hours per week) and does not report working part time in any week, then annual hours are estimated in the standard way as the product of weeks worked and hours usually worked per week. The same formula is used if an individual reports working part time throughout the year. However, individuals who usually work full time but work part time in some weeks (or who usually work part time but work full time in at least one week) are not asked for their hours during part-time (full-time) employment. To fill in this data gap for these workers, we use information on individuals who worked part time in the last week (not

year), but who usually work full time. We regress hours worked by such individuals in the last week on race, age, education, and usual hours/week and use the estimates to obtain a conditional expectation of the part-time hours/week of usually full-time workers. Annual hours are then calculated as the product of weeks worked full time and hours usually worked per week, plus weeks worked part time multiplied by the estimate of part-time hours. An analogous procedure is used to calculate the annual hours of individuals who usually work part time but work full time in at least one week.

6. These results are consistent with other recent studies; see Schor (1991) and Coleman and Pencavel (1993).
7. See Bound and Johnson (1992), Burtless (1990), Haveman and Buron (1994), Karoly (1992), Levy and Murnane (1992), and Moffitt (1990).
8. In predicting individual hourly wages, we first estimate annual selectivity corrected (Heckman 1976, 1979), hourly wage functions over all wage and salary nonwhite workers and white workers from data in the annual March CPS from 1976–1993. The independent variables are those exogenous human capital determinants of market productivity that are recorded in every CPS year. The race/year-specific coefficient estimates are used to predict each person's hourly wages based on his values for each of the attributes in the wage function. A more complete description of the procedure is found in Buron, Haveman, and O'Donnell (1995) and Haveman, Buron, and Bershadker (1997). The parameter estimates for the two race-specific wage functions for each year are available from the authors, as well as the probit equations that provide the basis for selectivity correction.
9. While labor market distortions may cause observed (and, hence, predicted) wages to be an imperfect measure of the productivity of an individual's work time, we accept these market values as the most appropriate weighting factor available for estimating the value of both earnings and potential earnings. We note that changes in labor market distortions over time will be reflected in the trend of aggregate measures of both earnings measures. For example, the presumed reduction in the influence of labor unions on wages (associated with the fall in union membership over the past two decades) could lead to a downward trend in both earnings and potential earnings due to a decrease in estimated wage rates. It should also be emphasized that the estimated wage rates used to weight actual and potential (2,080) work hours reflect the interaction of supply and demand factors in individual markets at a point in time. Hence, individual potential earnings estimates can only be aggregated to indicate the total, or per capita, value of potential earnings under the assumption that the structure of wage rates would not change in any important way if all males were to increase their annual work time to 2,080 hours, reflecting the full use of their human capital.
10. Given this convention, underutilization indicators could be calculated by comparing the actual hours that individuals work to the full capacity work hours norm of 2,080 hours. However, because we are interested in human capital utilization rather than labor hours utilization, we account for individual productivity as mea-

sured by the predicted wage rate in measuring both the earnings and the potential earnings components of FPE.

11. It should also be noted that in these cases, new jobs will be created in place of do-it-yourself activities.
12. Another indicator of the extent of labor underutilization is the percent of all working-age males who work less than the "full activity" norm, and hence record some level of FPE. We have studied this indicator of the "prevalence" of FPE and reported the results in Buron, Haveman, and O'Donnell (1995) and Haveman, Buron, and Bershadker (1997).
13. Aggregate earnings is the sum of the individual earnings of working-age males, which we described above as the product of an individual's actual annual hours of work and the individual-specific predicted wage rate. Dollar comparisons are in 1993 prices throughout the paper.
14. A regression of each of the three series in Table 2 on a time trend reveals average annual decreases of per capita earnings and potential earnings of $154 and $152, respectively, and an annual per capita increase in FPE of nearly $3. These findings indicate that the decrease in per capita earnings is the result of a decrease in both the level and realization of potential earnings.
15. In allocating foregone work hours to these seven reasons, we first split foregone work hours into hours per week and weeks worked deficits, and then allocated these separate components to the categories. For individuals who worked during the year, the unemployment reason was obtained from responses to a question regarding the number of weeks an individual was not working but was looking for work. In the survey, workers were then asked what they were doing for most of the remaining weeks of the year, with the following potential responses: illness/disability, taking care of home/family, retired, no work available, other. Any worker responding "no work available" had the value of these hours allocated to the discouraged worker effect. Other responses had these hours allocated as indicated. Individuals who did not work at all are also asked how many weeks they were in the labor force looking for work and these hours are attributed to the unemployment reason. These workers were then asked the reason for not working, with the following potential responses: illness/disability, taking care of home/family, could not find work, other. These responses were allocated in the same way as for workers. Individuals who report working part time for at least one week in the last year are asked for the main reason for doing so, with the following categories indicated: 1) could only find part-time, 2) wanted part-time, 3) slack work/material shortage, 4) other. In order to allocate foregone hours arising from part-time work to our categories, we supplemented the information on reason for working part time last year with information on the reason for working part time in the last week, and reason for working part year, and then proceeded to allocate responses similar to the procedures for workers. A more detailed description of these procedures is found in Buron, Haveman, and O'Donnell (1995) and in Haveman, Buron, and Bershadker (1997).

16. The average annual changes described hereafter were calculated from regressions of the relevant series on a time trends.
17. See Chirikos (1986) and Colvez and Blanchet (1981).
18. The attribution of FPE into "exogenous constraint" and "individual response" categories rests on a judgment over which people can disagree. For example, an individual may choose not to work, but may report illness (included in our "involuntary" category) in order to indicate a more acceptable reason for not working. The reason "other" is excluded from these estimates.
19. Tables describing the detailed subgroup patterns are available from the authors upon request.
20. The decreasing ratio of nonwhite to white potential earnings reflects the overall increase in wage inequality over the period. Because nonwhites are concentrated at the lower end of the education and skill distributions, increased wage disparity between the higher and lower end of these distributions is also reflected in increased wage and potential earnings disparities between racial groups over the 1975–1992 period. Recall that potential earnings is the product of the individual wage rate and a constant (2,080 hours).
21. It is possible that a growing fraction of older workers are switching from illness to retirement as the reason for not working.
22. Surprisingly, the dropout group reported that FPE due to retirement in 1975 ($770) was less than 20 percent of FPE due to illness ($4,616).

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6

The Growth of Income and Employment Inequality in Australian Cities

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Since the early 1970s, income inequality among individuals has been growing in most OECD countries. It has arisen from two sources: higher levels of unemployment, especially in Europe, and widening wage dispersions, particularly in the United States. Australia has also been subject to these trends, and the increasing inequality has led to a fast-growing research literature which documents the changes (Gregory 1993; Borland and Wilkins 1996; Saunders 1994).[1] The evidence seems to suggest that the change in inequality is less than in the United States and the United Kingdom.[2]

This chapter begins the process of analyzing changing income inequality on a spatial basis. It utilizes census data to emphasize changes in income and employment inequality within Australian cities over a period from 1976 to 1991. The data cover more than one-third of the Australian population. The analysis reveals a dramatic change in society. The shift in income inequality among individuals and families that has occurred over the 15 years from 1976 has been magnified on a spatial basis. Average household income has increased 23 percent in the 5 percent of neighborhoods with the highest socioeconomic status (SES), and fallen 23 percent in the 5 percent of neighborhoods from the lowest SES. These changes have been driven predominantly by employment changes. In 1976, employment activity of neighborhood residents was not related to the SES ranking of the neighborhood, but by 1991 that had changed. Employment in neighborhoods from the bottom 5 percent of neighborhoods ranked by SES status had fallen 37 percent.

The chapter is structured as follows. We begin by briefly describing the macroenvironment within which urban inequality has increased. The next section documents neighborhood changes according to the 1976 census and the 1991 census. We then demonstrate that the increased income inequality is being generated by employment shifts across neighborhoods. The following section conjectures as to the causes of these changes and offers some policy comments, and concluding remarks are contained in the final section.

THE MACROENVIRONMENT AND INCREASED NEIGHBORHOOD INEQUALITY

Some parts of the Australian labor market have performed well over the last two decades. The more successful features include a rapid growth of part-time jobs for women and young people. Some periods also exhibited strong aggregate employment growth, especially during 1983 to 1989 and 1993 to 1995. In addition, after 15 years of insignificant growth, average real wages have begun to increase again. Although there have been other good changes in the Australian labor market, poor outcomes dominate and four adverse features stand out in the period since 1976.

1) Employment opportunities for men and women seeking full-time work have not kept pace with population growth rates. A slight decrease in full-time male employment might be anticipated, as more men seek early retirement and younger men stay longer in education institutions. Since June 1976, however, the male full-time employment ratio has fallen 21 percent, which is far greater than what might have been expected (Figure 1). Unemployment among full-time male workers at May 1997 was 8.8 percent.

   Young women have also extended their involvement in education, but with the reduction in the birth rate, more divorces, postponement of marriage, and more women seeking careers in paid employment, it might be expected that full-time employment would increase. But at May 1997, the proportion of women employed full time was only 5.0 percent more than at August 1976. Unemployment among female full-time workers has increased from 4.3 percent in 1976 to 9.7 percent at May 1997.

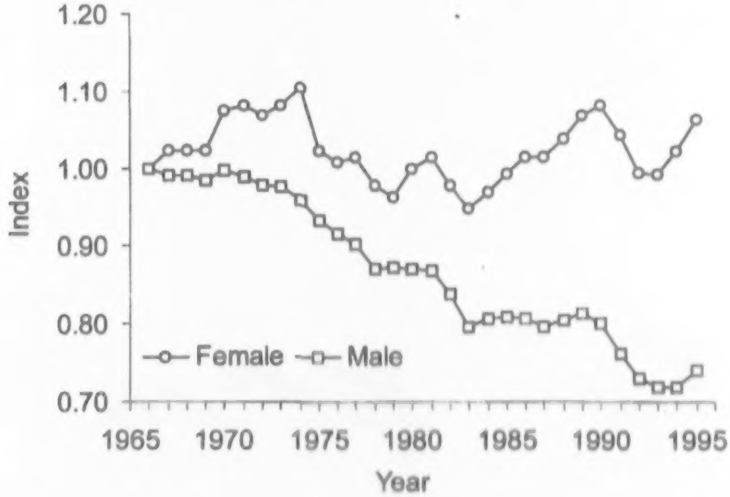
2) During each cycle over the last two decades, the number of welfare recipients, such as those who receive unemployment benefits, increased quickly and failed to return to previous levels during the recovery. This hysteresis effect suggests that much of the full-time employment reduction was involuntary.

3) The length of the unemployment spell has increased, and Australia has developed a long-term unemployment problem. In 1976, the average current spell length of unemployed persons was 17.5 weeks. By May 1997, the spell length had increased to 52.6 weeks.

4) There is a significant widening of the earnings distribution among those men who have been successful in obtaining full-time employment. Earnings inequality also increased among women (Gregory 1993).

These four adverse features suggest that economic and social inequality widened in Australia, and this is what most researchers find for most periods (Saunders 1994; Harding 1995). These studies analyze changes among individuals, and to a lesser extent changes among households or family units. It seemed to us that there should be spatial parallels within major cities where the rich and poor live in different locations.

Figure 1 Full-Time Employment/Population Indexes, 1966–1995 (1966=100)



NEIGHBORHOOD INCOME INEQUALITY CHANGES, 1976–1991

The Data

Australia has always had neighborhoods that are clearly demarcated by income and SES. Nevertheless, the undesirability and adverse effects of low income neighborhoods are not stamped on our national consciousness to the same extent that they are often stamped on the consciousness of citizens of other countries. United States citizens, for example, are very aware of the poverty of their inner cities and are well aware of the undesirable effects on residents (Wilson 1987; Case and Katz 1991).

The census is the only consistent database available to trace changes in neighborhood inequality over a significant period of time. There are four census collections that include income data that could be used to measure neighborhood changes. Each census—1976, 1981, 1986, and 1991—coincided with an economic recession. By some measures, the depth of the recessions are not too dissimilar, but it is noticeable that the unemployment rate is subject to an upward trend: 4.4, 5.6, 8.0, and 9.5 percent, respectively.[3] Because unemployment is higher at each successive date, we cannot use census data directly to analyze income distribution effects of economic cycles; therefore, we emphasize the trend from a comparison of 1976 with 1991.

To conduct the neighborhood analysis, the data are presented as group averages from collection districts (CDs), which are the smallest geographical area for which census data are available. CDs usually contain 200–300 dwellings that are delineated by easily identifiable boundaries. CDs tend to remain unaltered through time, and in our sample we exclude those which were subject to boundary changes and not comparable across the four censuses. The analysis is confined to CDs within major urban areas with populations of more than 100,000.[4] The panel consists of 9,483 CDs and about six million people in each of the four years. There are no other comparable data sets which allow such a rich analysis of the changing geographical distribution of economic variables. The results reported here are similar to those derived from post-code data, which, on average, groups CDs into population groups of about 4,500 (see Gregory and Hunter 1995).

Although the census provides by far the best data, they are not ideal. For example, income data are not available by source. Consequently, it is not possible to investigate directly the role of government welfare payments or other social services. There are no data on taxes paid. Another difficulty is that detailed geographic data are released as grouped means for specific variables, and it is not possible for us to reclassify the data in many ways that would improve our understanding.

The geographical analysis is based on CDs ranked by socioeconomic status (SES). We use the measure of SES calculated by the Australian Bureau of Statistics for 1986 (1990).[5] Each CD preserves its SES ranking over the 15 years. None of the results are affected by the choice of the census year on which the SES ranking is based.[6]

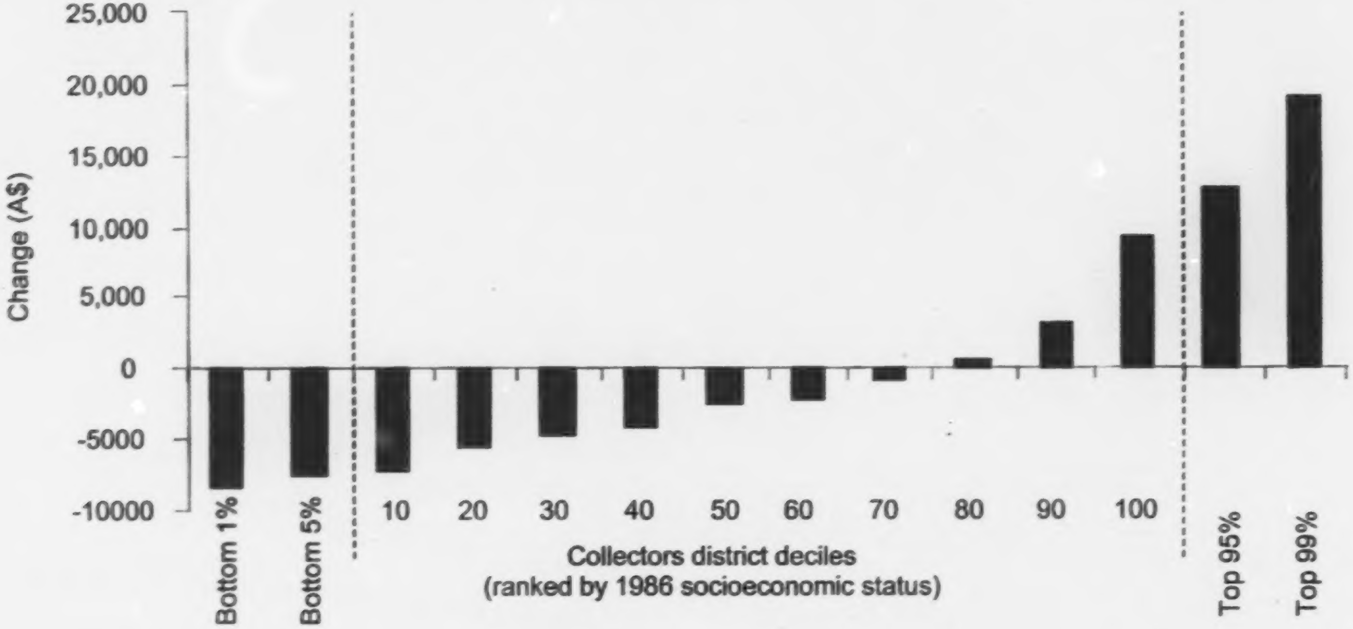
Neighborhoods and Household Income

We believe that income and employment gaps between our best and worst neighborhoods are not as great as the gaps in many major OECD cities. We also believe that Australia is not in danger of creating urban problems to the same degree as the United States.[7] However, we were surprised at the extent of the changes for the worse that have occurred since the mid 1970s.

We begin by discussing the marked change in the dispersion of annual household income across neighborhoods. In 1976, the ratio of the mean household income of CDs from the lowest to the highest 5 percent of SES areas was 60.4 percent. Within the space of 15 years, the ratio had fallen to 37.9 percent. Income distribution has become more unequal and is well beyond that which can be ascribed simply to changes in the structure of households. There is a significant increase in the geographic polarization of household income across Australia. The poor are increasingly living together in one set of neighborhoods and the rich in another. The economic gap is widening.

Figure 2 arranges CDs from low to high SES and enables us to identify the pattern of income change across CDs. The CDs are ordered on the basis of their 1986 SES rankings. The first two bars on the left measure the change in mean income over the 1976–1991 period

Figure 2 Change in Average Household Income, 1976–1991 (1991 A$)



for the 1 percent and 5 percent of CDs with the lowest SES. The last two bars on the right measure the change in mean income from the top 5 percent and 1 percent of CDs. All other bars refer to the change in annual household income averaged within each CD decile. Average income is in 1991 prices. Each decile includes approximately 500,000 adults.

As we move across the CDs from low to high SES areas, the pattern of income changes is quite smooth. For the bottom 70 percent of CDs, average household income has fallen in absolute terms and is lower in 1991 than in 1976. In areas of the highest SES, household income has increased markedly. In the top 5 percent of SES areas, household income has increased by $12,555 (23 percent). In the lowest 5 percent of areas, household income has fallen by $7,589 (23 percent). The income gap between the top and bottom 5 percent of CDs has almost doubled and has widened by $20,144 (92 percent).

This significant pattern indicates that the forces making for increased income inequality across households exert a strong and systematic neighborhood effect. These forces have either impacted upon individuals, according to the neighborhood in which they live, and/or there is a continual geographic sorting process at work so that households which lose income are moving to poor neighborhoods, and households which gain income are moving to high-income neighborhoods.

The narrow dispersion of neighborhood household income in 1976, and the increased inequality since then, are so notable that it is perhaps worth reemphasizing both facts by comparing household income from the top and bottom 1 percent of CDs ranked by SES. In 1976, the weekly income gap between average household from the bottom 1 percent of CDs and the average household in the median CD was not large (Table 1, column 1). An additional part-time job for nine hours per week at $12 per hour would close the gap of $116.

Facts such as this explain why most Australians believed that they lived in a fairly equal society in terms of income and employment opportunities. By 1991, however, an additional part-time job could still close the gap but it would need to extend to 19 hours per week, an increase of 10 hours. The bottom and median neighborhoods are drifting apart, and the gap has increased from $116 per week to $230 (1991 prices).

Table 1 The Neighborhood Household Income Gap

	Weekly difference 1991 A$			Change (%)		
Year	First percentile to median	Median to top percentile	Bottom to top percentile	First percentile to median	Median to top percentile	Bottom to top percentile
1976	116	442	558			
1981	175	430	625	51	–3	12
1986	227	620	844	30	44	35
1991	230	854	1084	1	38	28

The increased income necessary to move from the average household income in the median CD to the average household income of a neighborhood in the top 1 percent of CDs is larger. The additional income cannot be obtained from the usual part-time job. In 1976, the additional weekly income needed was $442, and by 1991 this had increased to $854 a week. This is not a small step. In 1976, the additional income might be earned from an additional job which paid a little less than average weekly earnings. In 1991, the extra annual income required was $44,408, an income level which far exceeds average weekly earnings.

The increase in income inequality across neighborhoods continued throughout the 15 years (Table 1, column 3) but the principal source of change differed. Between 1976 and 1981, increased inequality was generated by income falls in low SES neighborhoods. After 1981, the fall in income continued in low SES neighborhoods, but most of the increase in inequality was generated by income increases in high SES neighborhoods. The source of the increased inequality appears to have been shifting from large income falls in the low SES neighborhoods, relative to the median, to large increases in the high SES areas, relative to the median.

Neighborhoods and Male and Female Incomes

Figure 3 documents the change in the male mean annual income of CDs ranked by SES. Between 1976 and 1991, male annual income fell by $4,102 (1991 A$) in the 5 percent of CDs with the lowest SES. In

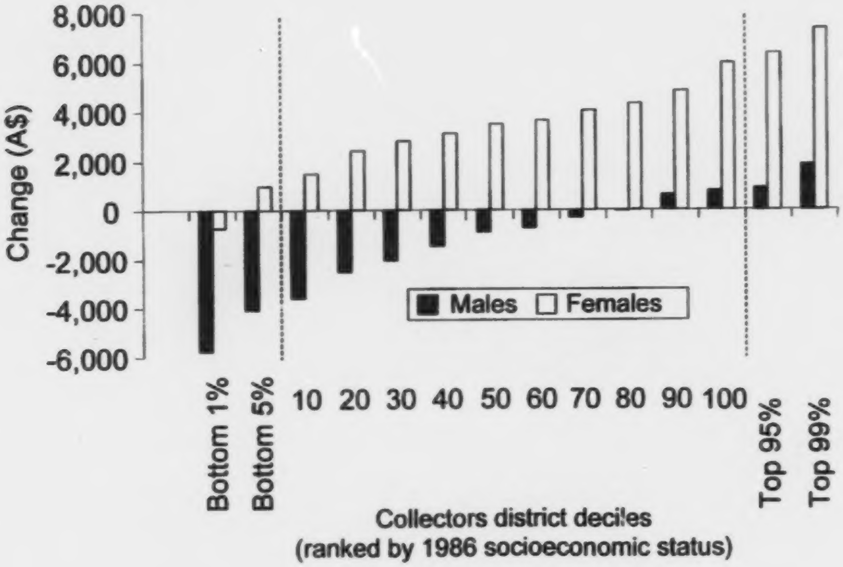


Figure 3 Change in Average Male and Female Income, 1976–1991 (1991 A$)

the top 5 percent of CDs, average male income increased by $916. As a result, the male mean income gap between CDs from the lowest and highest SES widened by $5,019.

It is noticeable that only 20 percent of CDs from the highest SES areas experienced male income growth over the 15 years. In 80 percent of neighborhoods there were real income falls.

The income changes for women also exhibit a smooth pattern across CDs (Figure 3). The mean annual income substantially increased in all but the lowest 1 percent of CDs, ranging from a fall of $726 for the 1 percent of CDs from the lowest SES areas to an increase of $6,321 for the 5 percent of CDs from the highest SES. Women's contribution to the income of a CD has offset the fall in male income, at least in part, in all but the lowest 1 percent of CDs.

Income distribution across neighborhoods has widened for both men and women. In 1976, the average male income in CDs from the lowest 5 percent of SES areas was 54.9 percent of the mean income in the highest five percent of SES areas. By 1991, this income ratio had fallen to 42.5 percent, a change not too dissimilar from the change in the household income ratio. The income level of women in the lowest to the highest 5 percent of CDs, ranked by SES, has fallen from 78.8 percent to 57.8 percent—once again, a change similar to that of the household income ratio.

EMPLOYMENT CHANGES AND THE INCREASE IN INCOME INEQUALITY ACROSS NEIGHBORHOODS

The Change in Male and Female Employment/Population Ratios

For most households, the principal source of income is employment. The relatively narrow income dispersion across neighborhoods in 1976 was generated by similar employment/population ratios across neighborhoods. For men, there was no systematic variation in employment/population ratios across CDs ranked by SES (Figure 4). For women, the employment/population ratio in 1976 was marginally less in low SES CDs, and the employment/population gap between the lowest and highest 5 percent of neighborhoods was small (Figure 5).

In 1976, irrespective of where they lived, Australians shared much the same commitment and access to employment. A social observer could walk across the best and worst parts of Australian urban areas, and although the probability of meeting someone who was employed differed by neighborhood, there was no systematic change by SES. Income inequality across neighborhoods ranked by SES was generated by different levels of income from all activities and not from differences in the proportions of the population employed.

By 1991, circumstances had changed dramatically. Australian employment growth between 1976 and 1991 had been very poor. Unemployment increased from 4.7 to 9.5 percent. The poor employment performance is evident in the neighborhood data. In all neighborhoods, the employment/population ratio for men had fallen—by 9 percent in CDs from the top 5 percent of SES neighborhoods, and by 37 percent in CDs from the lowest 5 percent of SES neighborhoods.

The pattern of employment change for women is similar, but the contrast across neighborhoods is greater. For the top half of neighborhoods, the proportion of women employed increased approximately

Figure 4 Male Average Employment/Population Ratio, 1976 and 1991

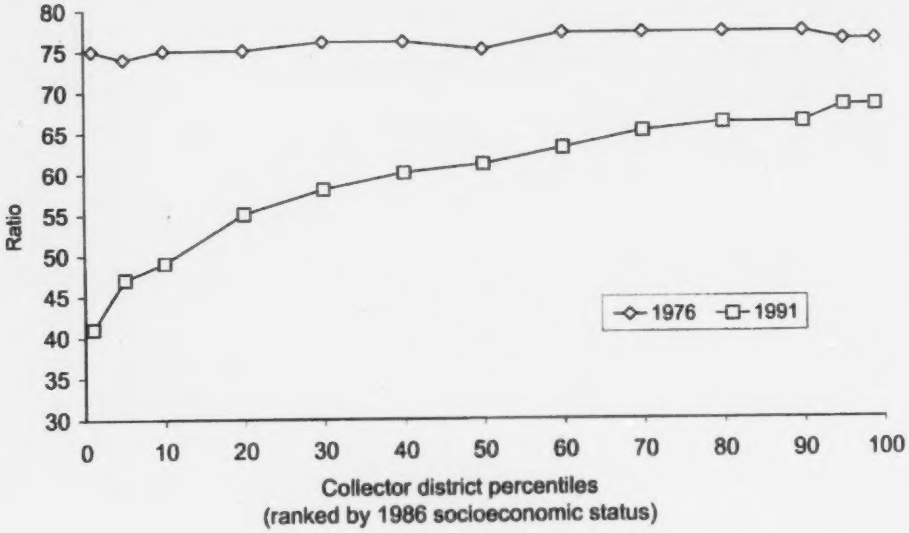
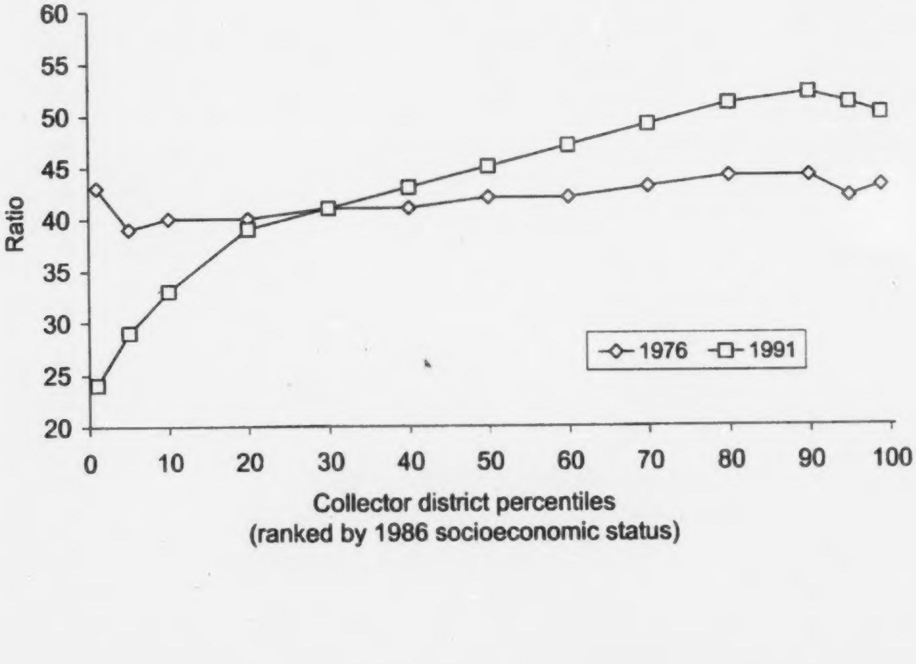


Figure 5 Female Average Employment/Population Ratio, 1976 and 1991



16.2 percent. The proportion fell 3.0 percent for the bottom half of neighborhoods, and 17.5 percent for the bottom decile. We are so used to seeing macrodata that indicate a rapid growth of part-time work for women and reading about women's increased labor force involvement, it is a shock to see that in 1991, and for half of Australian neighborhoods, the average proportion of women employed in the labor market is less than in 1976.

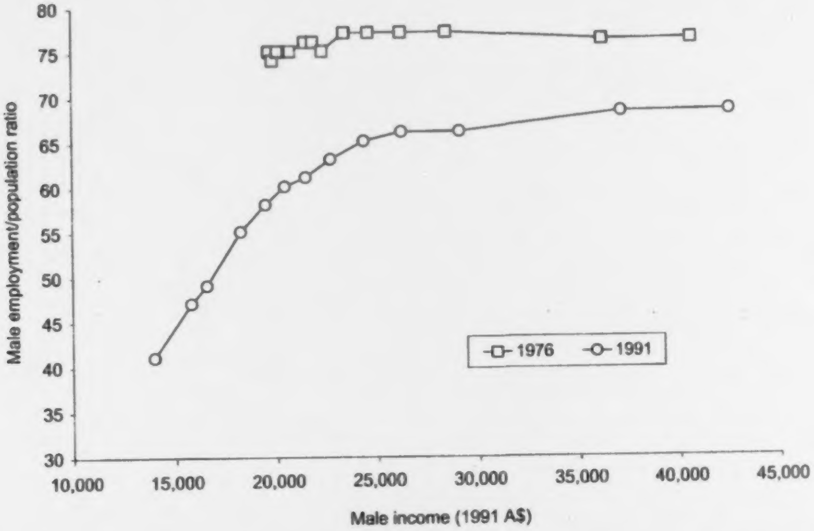
The growth in the women's employment/population ratio is concentrated in the high SES areas. By 1991 the probability that a woman would be employed if she lived in the top 5 percent of SES neighborhoods was 78 percent more than if she lived in the lowest 5 percent of SES areas.

It is apparent that employment/population ratios are now a major contributor to income variations across areas. For males, Australia has returned to the neighborhood employment patterns of the 1930s, with substantial pockets of non-employment. For women, however, the pattern is quite different (Gregory et al. 1987). In the 1930s, there was little variation of female employment/population ratio across neighborhoods ranked by SES. The pattern was much the same as in 1976. The loss of women's employment in low SES areas needs to be better understood.

The New Face of Australian Cities

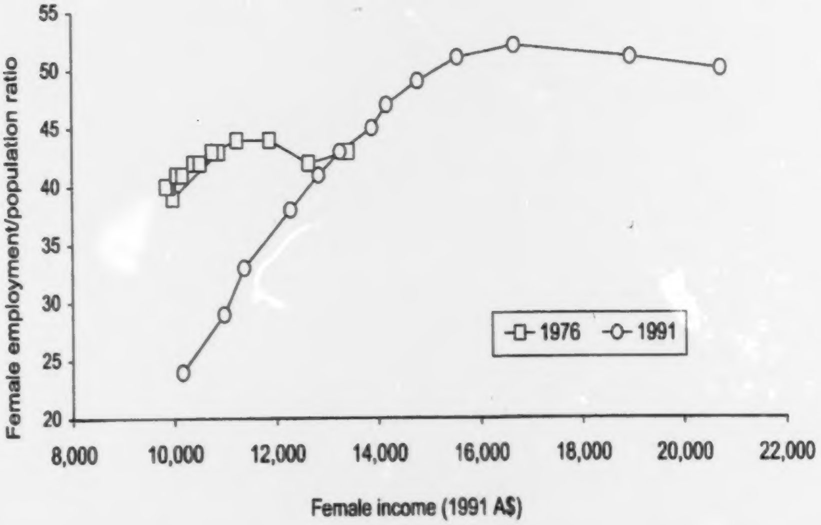
Neighborhoods in 1991 can be divided into two groups. For neighborhoods taken from the top 20 to 30 percent of CDs, ranked by SES, the employment/population ratio of men and women does not change significantly across neighborhoods, and there is no close relationship between employment level changes and income changes (Figures 6 and 7). Income dispersion within this group is related more closely to variations in wages and salaries and earnings from own business rather than variations in employment rates. For our social observer walking through the top 20 to 30 percent of neighborhoods, the level of employment has changed since 1976 but the pattern of employment across CDs has not. Employment/population ratios continue *not* to vary systematically across neighborhoods by SES and are not related to income changes.

Figure 6 Male Average Income and Male Employment/Population Ratios



NOTE: The points represent collector district percentiles, 1, 5, 95, and 99 and the average of male employment/population ratios and income, for deciles 0–10, 11–20, and 89–90. Data are taken from Figures 3 and 4.

Figure 7 Female Average Income and Female Employment/Population Ratios



NOTE: The points represent collector district percentiles, 1, 5, 95, and 99 and the average of female employment/population ratios and income, for deciles 0–10, 11–20, and 89–90. Data are taken from Figures 3 and 5.

For the remaining 70 to 80 percent of neighborhoods, employment rates now matter. The world has changed and there is now a clear association between employment changes and income changes. Within this group the translation of employment changes into income changes is similar for both men and women. On average, an increase in employment of 15 percentage points adds $2,300 to male income (see Figure 6) and $2,816 to female income of a neighborhood (see Figure 7).

The widening of the income distribution across neighborhoods is being driven by different influences at different ends of the income distribution. Employment is strongly associated with income in low-income neighborhoods but not in high-income neighborhoods.

Joblessness in low SES areas begins with teenagers (Figure 8). In 1991, the employment rate of teenagers in low SES areas is 80 percent of that of high SES areas, even though most teenagers in high-status areas are attending an education institution. Within the age group of 20–24 years, the employment rate of the bottom 5 percent of CDs has fallen to 63 percent of that of the top 5 percent of CDs, and it remains there until the age group of 45–54 years, where the employment rate falls further.

Figure 8 1991 Employment/Population Ratio for All Persons, by age, in Lowest and Highest 5% of SES Areas

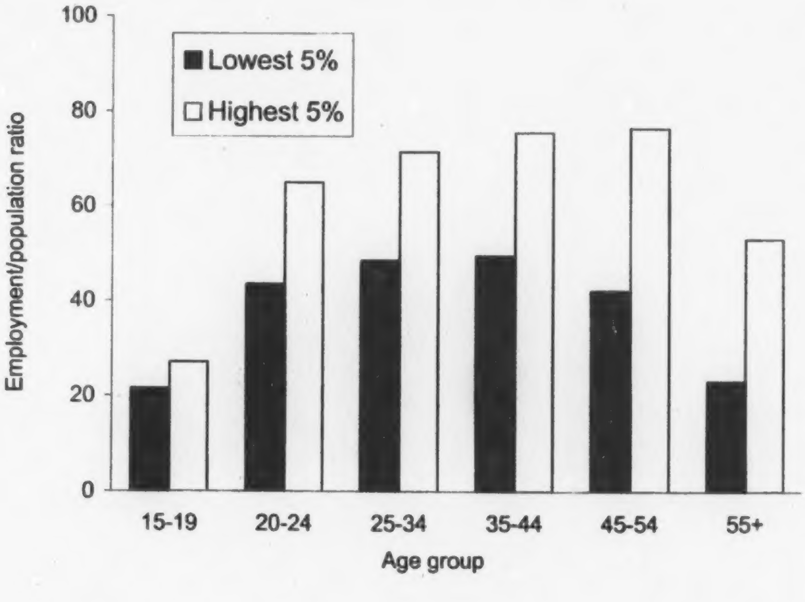


Table 2 Change in Employment and Real Income in Public Housing Neighborhoods and Other Neighborhoods in the Bottom 10% of SES Rankings, 1976–1991

	Public housing[a] (%)	Neighborhoods no public housing (%)	All (%)
Real Income			
Male	–29	–13	–18
Female	–2	17	13
Personal	–19	–1	–7
Household	–34	–12	–21
Employment			
Male	–42	–24	–29
Female	–30	–5	–11
Total	–37	–15	–22

[a] Public housing neighborhoods: 50 percent or more of the neighborhood population residing in public housing. There are 207 public housing neighborhoods in the sample.

The pattern is the same for men and women. It is remarkable that in 5 percent of CDs from the low SES areas, almost one-half of the men 25–44 years are not engaged in employment.

CONJECTURES ON CAUSES OF INCREASED URBAN INEQUALITY AND POSSIBLE POLICY RESPONSES

Although we are concerned about the rapid growth in income inequality across neighborhoods, it is nevertheless true that there is no "right" degree of urban inequality. Nor is it clear that policy can efficiently and effectively achieve the urban inequality we might prefer. In the past, Australia has not placed high priority on policies specifically directed toward reducing urban inequality, and our experience of policy effectiveness in this area is limited. Policy has been more concerned with income distribution and unemployment among individuals. However, what can be done if we are dissatisfied with a situation where, in

1991, male unemployment is as high as 35 percent in many neighborhoods? How can we return to something approaching the distribution of neighborhood income in 1976? It is not possible to answer these questions without some understanding of the underlying causes of urban inequality growth.

Public Housing Policy

Increased neighborhood inequality and public housing policy have been closely intertwined. Approximately 5 percent of the Australian population live in public housing, which is usually found in areas of low SES. As unemployment has increased, access to public housing has become more focused on the poor and economically disadvantaged, and the economic circumstances of the typical public housing resident has changed considerably for the worse. Table 2 is confined to CDs located in the bottom 10 percent of SES neighborhoods. It shows over the 1976–1991 period the income change in public housing areas, which we define as CDs where the proportion of the population in public housing exceeds 50 percent. It is evident that public housing neighborhoods have done much worse than other low SES neighborhoods. Over this period, the average real income of a male in a public housing neighborhood fell 29 percent. The average real income of a male in nonpublic housing neighborhoods fell 13 percent. For women, average real income fell in public housing neighborhoods by 2 percent and increased by 17 percent in other neighborhoods.

Employment changes are also large and negative in public housing neighborhoods. Employment of men and women fell 42 percent and 30 percent, respectively. In nonpublic housing neighborhoods, employment fell 24 percent for men and 5 percent for women.

For most of the period, public housing policy increasingly grouped low-income people together and contributed to the falling income in low SES neighborhoods, but that is only a part of the story. In low-income neighborhoods without public housing, there are also substantial but lower employment and income falls.

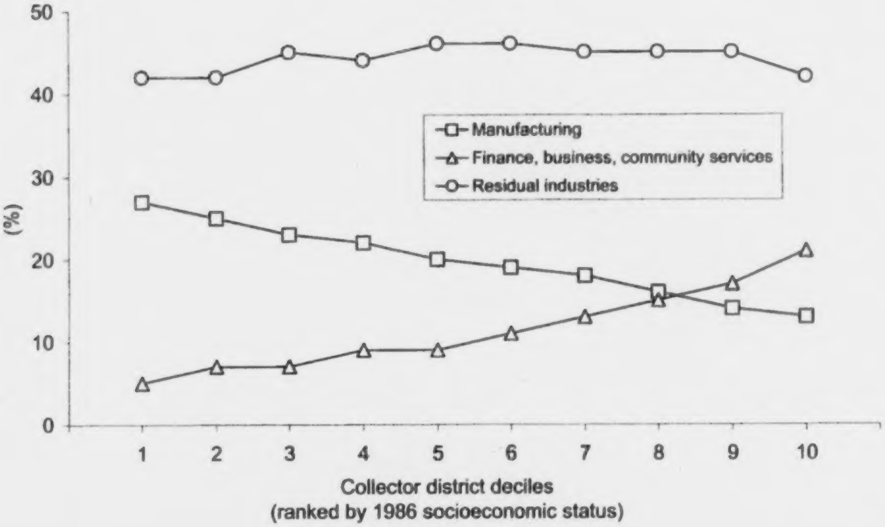
As falls in employment and income of public housing neighborhoods have been so substantial, these neighborhoods have come to be seen as areas of social deprivation that are creating environments of poverty from which public housing tenants are finding it hard to escape. Toward the end of the period, therefore, public housing policy began to change and slowly attempted to disperse tenants more widely in the community.

Manufacturing Decline

Another important influence generating increased urban inequality seems to be the rapid decline in manufacturing employment. A glance at the 1976 census data is sufficient to indicate that the rapid decline in manufacturing employment has generated important spatial shocks within cities. To illustrate this, we divide industry of employment into 12 two-digit Australian Standard Industrial Classification categories and focus on the male labor force. Similar considerations apply to the female labor force.

Figure 9 plots the proportion of men over 15 years of age who were employed in manufacturing within each CD in 1976. The horizontal axis orders CDs by their 1986 SES rankings. Individuals are classified by area of residence and not by location of employment.

Figure 9 Proportion of Males Employed by Industry Group, 1976



NOTE: The points represent the average of collector district deciles.

There is a distinctive pattern. In CDs from the bottom SES decile, 27 percent of all males over 15 years of age were employed in manufacturing. As the SES of the area increases, the manufacturing employment proportion falls declining to 13 percent in areas of high SES. Figure 9 also includes 1976 male employment in finance/business and community services. Five percent of men over 15 years of age from the bottom SES decile are employed in these industries. In areas of high SES, these two industries employ 21 percent of all men. Employment in the other nine industries, which we label the residual category, exhibit no noticeable and systematic pattern across SES areas.

Between 1976 and 1991 there was a large negative macroshock to manufacturing, as male manufacturing employment, as a proportion of the male population over 15 years of age, fell 37 percent. Labor market changes in other industries did not help the employment adjustment that was required. Employment in the residual industry category, as a proportion of men over 15 years, fell 14 percent and did not provide opportunities for net job growth. The pattern of decline was much the same irrespective of the SES ranking of the neighborhood. The only significant source of male employment increase, 29 percent, was in finance/business and community services, where employment change favored high SES areas. The net result is that the male manufacturing employment loss in low SES areas was not offset. Men who live in low SES areas were not able to make employment inroads into other industries.

It is perhaps not surprising that the job loss was spread unevenly across CDs and fell disproportionately in areas where manufacturing employees live; this is to be expected given the initial employment pattern. The interesting point is the spatial nature of the persistence of joblessness. What could be the mechanisms generating these outcomes? At this stage we do not know. One possibility is the following: suppose, as a rough approximation, that finance/business and community services tend to locate in the city center or in local shopping and business areas that are easily accessible to all potential employees. Transport routes are focused on these locations. Industries in the residual category are spread randomly throughout the community and therefore jobs are easily accessible as well. Factories, however, are clustered and spread unevenly throughout the city but are close to low SES areas where the majority of their workers live.

If this description is broadly correct, when factories close they create local areas of unemployment. There are residual industry jobs nearby but the total number is contracting. The expanding finance/ business and community services sectors are located in areas which involve greater transport costs and, in addition, the job growth in this sector has not been sufficient to absorb manufacturing job losses. The persistence of the geographical dispersion of unemployment arises because of structural changes across industries, the geographic location of the lost jobs, and increased transport costs to gain access to new jobs.

The persistence elements of the analysis can be reinforced by other changes that are occurring in the economy. Suppose that at the same time factories are closing, welfare payments for those who cannot find employment are increasing in real terms, transport costs are increasing in response to the movements toward less subsidies, and real wages are falling among low-paid workers. Lower real wages offered to those at the bottom of the wage distribution may encourage some people to remain in a job-loss area and live on unemployment benefits—which in Australia have no time limit—rather than to accept employment at lower wages and incur higher transport costs. Furthermore, if house prices and rents respond to the lack of work in particular parts of the city, the effects of regional specific shocks will be increased. A wider variance of rents, reflecting a change in the ease of finding employment from each geographic base, may encourage people to stay unemployed and pay low rents rather than move to a high-rent area, give up unemployment benefits, and accept a low-paying job.

If mechanisms similar to this are generating spatial persistence of unemployment in areas where manufacturing workers used to live, a number of important points follow. First, the unemployment problem cannot be solved by macropolicies that do not create a job bias toward those areas. Second, trends in the key variables—increased transport costs, increased welfare payments relative to wages at the bottom of the wage distribution, and a falling proportion of employment in manufacturing—seem unlikely to be significantly reversed. Hence, in the absence of some intervention, unemployment may continue to persist on a geographical basis.

General Macro Influences

Increasing inequality may also be the result of major structural problems in the macroeconomy—such as emerging inflation or balance of payment difficulties—that lead to restrictive macropolicies and insufficient job creation. Irrespective of the initial nature of the adverse macroemployment shocks, those with more skills find jobs quickly and displace the least skilled, who eventually become unemployed. The unemployed gradually sort themselves geographically so that eventually more and more of the jobless live in depressed areas where the rents are lowest.

This explanation would suggest that the correlation between the decline in manufacturing employment and job loss by area is of no special significance. When the economy recovers and sufficient jobs are created, the updraft draws individuals from low SES areas back into employment and back into higher income levels.

One piece of evidence that might support this view is that, according to census data, approximately 40 percent of males living in a CD were not resident there five years earlier. This mobility raises the possibility that males who lose their jobs in manufacturing leave the CD and are replaced by others who are unemployed but not necessarily as a result of manufacturing decline. To confirm this we still need to know the SES status of the areas where individuals move to and come from, but the census does not provide that information at the detailed level at which this analysis is conducted. This is an important piece of missing data.

If individuals move a small distance to an area similar to the one they left, that might be considered as being the same as no mobility. The economic and social environment of those that moved, and their propensity for obtaining employment, may not have changed. If individuals leave to find jobs in better areas, we need to ask what it is about the low status areas where manufacturing employees used to live that leads to unemployment persistence.

It is unlikely that the unemployment increase since 1976 can be attributed to only one cause and be fully explained by a simple model. The facts, however, suggest that there are significant regional shocks within cities and these shocks may lead to unemployment persistence. If so, then a new research agenda is needed—one which combines the textbook macroanalysis of unemployment with regional specific shocks and persistence.

Relative Wage Flexibility

Some may argue that the best policy response is to increase labor market flexibility so that wages can fall in low SES areas and thereby create jobs. It is not known how much wages might need to fall, but to increase employment of the bottom 5 percent of SES areas back to 1976 levels, relative to high SES areas, there would need to be at least a 44 percent increase in male employment and a 70 percent increase in female employment. It appears likely, therefore, that a substantial wage fall might be required. This raises a number of problems. First, it takes time to create jobs so that the short-run wage fall might be substantial—so substantial, in fact, that individuals may prefer not to work and be supported by unemployment benefits and other welfare payments, and perhaps a range of black economy activities.

If wage reductions were to occur, yet low employment rates persisted in low SES areas, it might be expected that governments would eventually react and reduce benefit levels, relative to low wages. Labor market–related benefits are the main source of income for most individuals in low SES areas, and any reduction must inevitably increase poverty and widen income distribution further. It is obvious why governments and communities are reluctant to go down the path of substantial reductions in wages and benefits, and why it is often suggested that it might be better to try and increase the employability of individuals in low SES areas rather than reduce their potential wage.

Education Policy

Many countries, including Australia, have attempted to use an expansion of education and skill training to offset growing income inequality and unemployment among the low paid. Students in Australia have been offered means-tested living allowances for high school and tertiary education and interest-free loans to pay university fees. Tertiary and high school places have increased substantially. Indeed, over the last decade and a half, Australia has embarked upon one of the most ambitious education programs in the OECD.

This education expansion has had a large impact on the average neighborhood from areas of median SES. Between 1976 and 1991, the proportion of the population with degrees increased from 3.7 to 14.7 percent and the proportion of the population without qualifications fell from 66 to 45 percent. Yet despite this large increase in education of the potential workforce, male unemployment in median neighborhoods has risen from 4.4 to 13.0 percent. In addition, average income per adult has risen by less than one-half of 1 percent per year.

Income and employment outcomes may have been worse without education increases, but it appears, nevertheless, that increased education levels have not been sufficient to offset significant employment losses or to generate significant income increases for the median neighborhood. Education and skill training may primarily determine who gets jobs and may have very little influence on the number of jobs available or average rates of pay.

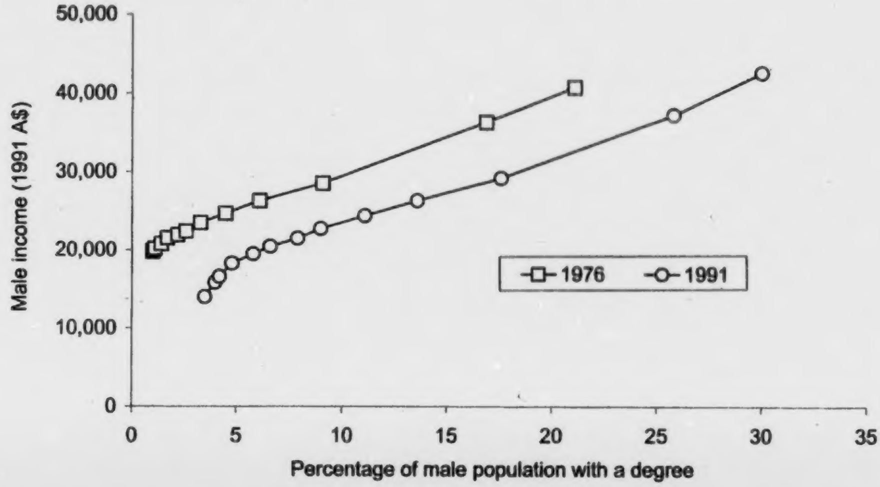
A similar sober assessment also appears inescapable from a comparison of the changing interrelationship between education levels and income inequality among neighborhoods. Various measures of the education levels of a neighborhood's residents are highly correlated, and for our analysis we use the proportion of residents 15 years and over with a degree.

In 1976, there was a strong positive association between the average education level of a neighborhood and the income of its residents. On average, a 1-percentage-point increase in numbers of men holding degrees was associated with additional neighborhood income of $1,000 (Figure 10). For women, the relationship was $500 for each additional percentage-point increase in the proportion of the female population with degrees (Figure 11). Among neighborhoods, as among individuals, higher education brings higher income.

It is noticeable, however, that in 1976 there is no systematic relationship between employment-population ratios and education for either men or women. More education is associated with more income but not because employment is increased. This is a restatement of the fact that in 1976, employment opportunities were distributed equally across neighborhoods ranked by SES.

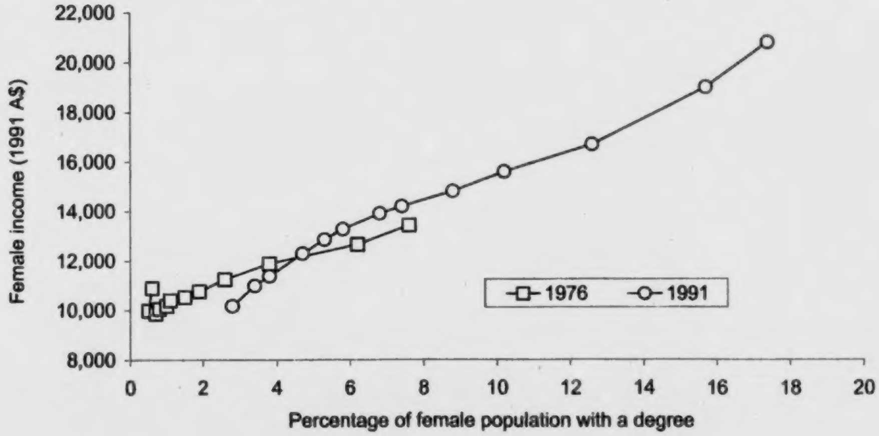
By 1991, the relationships have changed a great deal. For men, more education is still positively associated with more income, but the relationship has shifted so that for any given proportion of the popula-

Figure 10 Male Income and Proportion of Male Population with a Degree



NOTE: The points represent collector district percentiles, 1, 5, 95, and 99 and the average of male income and the proportion of the male population with a degree, for deciles 0-10, 11-20 89-90.

Figure 11 Female Income and Proportion of Female Population with a Degree



NOTE: The points represent collector district percentiles, 1, 5, 95 and 99 and the average of female income and the proportion of the female population with a degree, for deciles 0–10, 11–20 to 89–90.

tion with degrees, the annual income level has fallen by about $8,000. If the employment–education relationship can be thought of as a causal one, then in order to achieve the same level of male income as in 1976, a neighborhood needs to achieve a higher education level. Consider a neighborhood from a low SES area: to maintain male income, this neighborhood needed to increase the proportion of its male population with degrees by 6 percentage points between 1976 and 1991. The actual increase was 2.5 percentage points, hence the fall in male income. In high SES areas, the increase needed in the proportion with degrees was around 8 percentage points. The actual increase was 9 percentage points, hence the increase in male income.

This shift in the education–income relationship is very important. On the basis of the 1976 relationship between the incidence of degrees and the income of a neighborhood, the increased education attainment of the average neighborhood within the bottom five percent of CDs should have brought about an income increase of $3,500. In fact, there has been a fall of $6,000. The $9,500 gap clearly illustrates the importance of the change.

The principal source of the shift in the male education–income relationship is a shift in the employment–education relationship. For neighborhoods from the bottom 70 percent of SES areas, the education–employment relationship has moved down but, in addition, there is now a strong neighborhood relationship between less neighborhood education and less neighborhood employment—a relationship that did not exist in 1976. The lower the male education level of a neighborhood, the lower the male employment-population ratio. Education not only affects income, as it always has, but now it also affects the employment-population ratio. Poor neighborhoods are now twice disadvantaged by low education levels.

For neighborhoods from the top 30 percent of SES areas, further education does not bring further employment. Nothing has changed for these neighborhoods with respect to changes in education and changes in employment. But the education–employment relationship has also shifted downward, so at each neighborhood education level there is 15 percentage points less employment.

Labor market changes for women are similar to those for men in all but one respect—the education–income relationship has changed little since 1976 except in areas of low SES, where additional degrees among residents have not brought neighborhood income increases. But, unlike the male relationship, the large increase in women's income across all but the low SES areas is associated with the large increase in education. There has been no systematic shift down in the employment–income curve as in the male labor market.

There is a clear dichotomy between neighborhoods. For the top 30 percent of SES areas, income has fallen for each education level for men but increased for women. The relationship between changes in income and changes in education, however, has not shifted for this group.

For the remaining 70 percent of neighborhoods, the lower the education level the greater the income fall. Employment and education are now associated and hence there is less income at each education level.

To conclude, we look at the change in the distribution of education levels across neighborhoods to assess the general impact of the large increases in education levels of the potential workforce. In 1976, 10 percent of all residents 15 years of age and over who resided in CDs from the top 5 percent SES possessed degrees; now the proportion is 20 percent. In the lowest 5 percent of CDs, the proportion of the population with degrees has increased from 0.5 percent to 3 percent. The absolute gap in the degree distribution between areas has widened, and the increased incidence of degree qualifications has been disproportionately concentrated in CDs with high SES. Neighborhoods have not become more equal. For every 10 new degree holders in the top 5 percent of CDs, there has been an additional 3 in low SES areas. A similar pattern is evident if different measures of education are used.

Areas of low employment and low income have not been untouched by the expansion of education. Education levels have increased across all neighborhoods, but two major problems have emerged. First, the increase in education in absolute terms has been greater in high SES areas so that inequality has increased. Second, the relationship between employment and education levels has shifted in low SES areas such that a given level of education now delivers much less income and the move to a more disadvantageous relationship has dominated the improvement in the education level.[8]

It is a well-known finding in education research that school outcomes are related to the education level of the parents of the students who attend the school. The widening parental education gap across neighborhoods suggests that in order to expand the education opportunities for young Australians, special attention should be given to education policies directed toward schools in low SES areas.

CONCLUSION

Since the early 1970s, the Australian economy has had a major problem with job creation. According to the census, the proportion of men aged 15–64 employed in a median neighborhood is 19 percent less than in 1976. The proportion of women employed is 1 percent more. The shortage of jobs has not been rationed evenly throughout our society. Job loss and income falls are concentrated in low SES neighborhoods, and job growth and income rises are concentrated in neighborhoods of high SES.

Between 1976 and 1991, the lowest 1 percent of neighborhoods, based on a 1986 SES ranking, have lost 45 percent of their employment and 23 percent of their household income, and male unemployment has increased from 6.4 to 28.1 percent. The contrast with areas of high SES is marked: in the highest SES areas, employment has fallen marginally, household income has increased by 31 percent, and male unemployment has increased, but only to 4.8 percent. The proportion of women employed in high SES areas now exceeds by 20 percent the proportion of men employed in low SES areas.

To lose employment and suffer significant income losses are bad outcomes for anyone, but does it matter that these undesirable outcomes increasingly possess a spatial component? It is sometimes suggested that it does not and that nothing is gained by knowing that it is people who live in poor neighborhoods who are increasingly not at work, that part-time jobs are going to young people and women who live in high SES neighborhoods and that income is rising in the best SES neighborhoods but falling in poor neighborhoods. Our intuition suggests that neighborhoods do matter.[9] It seems likely that the greater the economic polarization within our cities the less equal are the opportunities for young people and the more likely that bad neighborhood pathologies will emerge. But there is not widespread agreement on these matters among Australian researchers.

But what should be done? It is not easy to know. There has not been a strong Australian tradition of thinking about economic policy and neighborhoods and it is not always easy to move from thought patterns that revolve around individuals or the macroeconomy to thought patterns that stress geography. There is also not widespread agreement as yet whether the growth of inequality across areas is just the natural outcome of more inequality among individuals, the impact of concentration of those individuals within a location, or whether the nature of the geographical areas is contributing to the inequality growth.

There is always more to be done. We do not know enough about social and geographical mobility, the role of job-finding networks and changing income, and employment opportunities over the lifetimes of people who live in poor neighborhoods.

Notes

R.G. Gregory is Professor of Economics and Head of the Economics Program, Research School of Social Sciences, Australian National University; Dr. Boyd Hunter is Research Fellow, Centre for Aboriginal Economic Policy Research, Australian National University. Cathy Baird and Eva Klug provided invaluable assistance in preparing the paper. This research was funded in part by the former Department of Housing and Regional Development. All dollar values in this chapter are in Australian dollars.

1. There is no consensus, however, as to the source of these large changes. They seem to be related to shifts in labor demand away from men and toward women workers and away from the unskilled towards those with higher education levels. There are some areas of agreement among researchers as to what is *not* driving the increased inequality. It does not seem to be the case that inequality among individuals is being driven primarily by the decline in manufacturing, the growth of trade with Asia, or immigrant flows of low skilled labor. We are more agnostic.
2. While there is a general consensus in this research that market incomes have become more unequal, the situation with respect to other measures of income is less clear. Government intervention in Australia has a strong equalizing component. Harding (1995) has estimated that the ratio of market incomes between the top and bottom 20 percent of the Australian population is 12.5:1. This reduces to 4.9:1 once transfer payments are taken into account, 3.8:1 after income tax, and 2.9:1 after government expenditure on services such as education and health.
3. Unemployment at August each census year taken from the Labour Force Survey.
4. CDs were omitted from the panel if the total population was less than 50 to avoid the sampling error deliberately introduced by the Australian Bureau of Statistics (ABS) to protect the confidentiality of persons in the neighborhood. In each suc-

cessive census, new CDs are added and in some circumstances the boundaries of CDs are changed. Our sample is a fixed number of CDs with unchanging boundaries that are to be found in each census plus a small number where the CD may have been divided into two. We begin with a list of CDs from the 1986 Census, and if there was more than one CD that corresponded to the 1986 CD, the first was taken to be representative of the 1986 CD.

5. As a measure of socioeconomic status, we use the Urban and Rural Indexes of Relative Advantage, published by the Australian Bureau of Statistics (1990). The Indexes are calculated by the application of principal components. The relevant variables include data such as family income greater than $50,000, the proportion of CD residents with degrees, the occupational distribution of the employed workforce, and the number of bedrooms per household.
6. In 1976, our sample of CDs represented 69 percent of all Australian CDs. By 1991, the sample had fallen to 52 percent. The average employment and income levels in new CDs in outer suburbs are a little higher than our sample means, so our sample understates slightly the growth in average employment and income over the 1976–1991 period, but our estimates of increased inequality are not affected (Hunter 1996).
7. The poverty of the U.S. ghettos is compounded by the concentration of disadvantaged Americans of African descent (see Wilson 1987). Another contributing factor is the U.S. Federal system that places emphasis on local taxes as a revenue source. The Australian federal system, in contrast, is a force for equalizing income and government services across neighborhoods.
8. The very large expansion of education must have affected the quality of education, and that may well have locational aspects. There is evidence indicating high failure rates in areas of low SES.
9. In a recent U.K. study, Gregg and Wadsworth (1994) show that the most successful method utilized by unemployed males to find a job is through friends and contacts. The utilization rate of this method is not the highest but it has the highest success rate. Among males, one-third of jobs are found this way; among women, one-quarter. Montgomery (1991) estimates that 50 percent of all workers currently employed in the U.S. found their jobs through friends and relatives.

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Part III

7

Supply of Hours per Day and Days per Week—Evidence from the Canadian Labour Market Activity Survey

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A significant amount of research has been conducted on the determinants of the supply of labor. In these studies the quantity of labor supplied is usually counted as the number of hours supplied per year or per week, largely owing to the fact that most labor force surveys do not disaggregate work hours more finely than the weekly level.[1] The number of hours per week, of course, is simply the product of days worked per week and hours worked per day, assuming both remain constant. Still, there is reason to believe that different workers desire to work a different number of days per week and hours per day, even though the number of weekly hours that each wishes to work may remain constant. For example, many individuals in the nursing profession regularly work three 12-hour shifts per week. This is not necessarily in response to the lack of options, but rather because they select into an occupation that offers a variety of days/hours combinations. In such cases, the use of weekly, monthly, or yearly hourly aggregates may mask a number of interesting characteristics of labor supply. For one, the fixed costs of supplying labor may differ depending on the unit of analysis. It is well-known that daily costs of work in terms of child care expenses, commuting costs, etc., may affect daily labor supply decisions. There may also exist hourly costs of employment which could likewise influence this dimension of labor supply. Furthermore, employer constraints on the hours and days that one is able to work could limit the optimal days/hours combination from the employee's point of view. To the extent that these constraints exist, optimal days/hours pairs may only become available as employees sort into new positions that offer a more desirable package.

A better understanding of the desires of individuals in choosing their hours/days combinations could ultimately lead employers to offer competitive weekly work schedules to employees, thus reducing turnover and absenteeism rates as well as related costs. For governments, such information may be useful in planning for future infrastructure projects, or how best to target day care subsidies.

This chapter will investigate in some detail the days per week and hours per day decisions of workers. The use of a unique data set allows us to decompose the usual weekly hours aggregate into daily hours and weekly days. The relevant literature will be discussed in the next section. The third section presents the data to be used in subsequent analyses. A preliminary look at the patterns of weekly working times for males and females, both paid employees and self-employed workers, is covered in the next section. A simple econometric model of supply of hours and days is the topic of the following section. This largely serves as a check on the data and will allow us to investigate further some of the pertinent determinants of the hours/days labor supply decision. The next section presents and estimates a simple model of job change behavior. Since individuals may be constrained from working their desired days and hours at any one job, they may change jobs in response to these constraints. Following that is a more detailed look at the actual hours and days changes of job-changers. The final section concludes and offers some areas for potentially fruitful future research.

Our results show that individuals tend to be clustered around a standard five-day, eight hours per day workweek, with men exhibiting much less flexibility around these norms than women. The self-employed, regardless of gender, are much less likely to work standard hours and days compared with those engaged in paid employment. It is well-known that women supply less labor when they have young children. But our evidence also shows that women with young children supply less labor, in terms of both hours and days, than those without young children, although the percentage drop in days is larger. This is not well-known and suggests that the costs of childrearing are borne on a daily rather than an hourly basis. In other words, women with young children find it more cost-effective to reduce days when reducing weekly hours.

We also discovered that job-changers desire flexibility in their weekly schedules. Not only do they display a larger variance in hours and days at their initial jobs compared with those who did not change jobs, but this variance increases further as they move into their new positions. This suggests that employees may be constrained within jobs from attaining their desired hours/days combination.

PREVIOUS RESEARCH

Previous research has addressed a variety of related time aggregation problems. Hanoch (1980a,b) distinguished between the hours per week and weeks per year decision in a reservation wage model of female labor supply. Blank (1988) built on this model to allow for simultaneity of the hours and weeks decision. She also allowed for discontinuities in the labor supply decision that can occur as a result of fixed costs of employment or if workers are constrained by firms who will only allow a minimum number of hours per week and/or weeks per year. She concluded that the evidence provides support for the theory that female heads face either significant fixed costs of employment or structural barriers to low levels of yearly weeks or weekly hours of work.[2] Also, decisions regarding hours of work per week and weeks of work per year were made independently, albeit simultaneously. The lesson is that using aggregated annual hours in many analyses may be inappropriate because the variable lacks the necessary detail.

The recent literature on Canadian labor supply has also analyzed the changes in hours worked over time, usually at the aggregate of annual or weekly hours worked, and often in the context of an explanation for earnings polarization. Morissette, Myles, and Picot (1993) have shown that the 1980s experienced a widening in the distribution of annual hours worked between workers. Morissette and Sunter (1994) and Morissette (1995) showed that the distribution of weekly hours also widened during the 1980s; fewer individuals worked 35–40 hour weeks, while the fraction working either shorter or longer hours rose. Other research has addressed the increase in multiple job holdings and part-time work (Krahn 1995; Logan 1994; Pold 1994, 1995). One of the lessons of this research is that aggregate measures of employment, such as annual hours worked, tend to obscure the fundamental underlying changes in the labor market over time, even though aggregate means may show only modest changes.

Just as aggregated annual hours hide important details, it is plausible that weekly hours may also be an inappropriate unit of analysis because the choice of days per week and hours per day could be related and simultaneous decisions on the part of workers. In addition, there can be fixed costs per day of work and even costs per hour of work. Aggregation of work hours into hours per year and hours per week does not allow us to investigate the complexity of the workers' decisions.

The costs of child care are frequently used in estimating the probability of female labor force participation. Many of these studies are nicely summarized in Cleveland et al. (1996). Blau and Robins (1988) and Ribar (1992), for example, found that child care costs had a negative effect on female labor force participation decisions. Cleveland et al. (1996) arrived at similar results using Canadian data. Generally, such empirical work is supportive of economic theory in that higher costs of child care lead to lower female labor force participation rates. What these studies have in common is the use of female participation as the dimension of labor supply analyzed. One exception to this is the study by Michalopoulos et al. (1992), which used hours supplied as the unit of analysis. They discovered that reduced child care tax credits resulted in a reduction in hours for women currently employed. These studies, however, did not address the impact of child care costs on the supply of hours and days. In another example of the importance of fixed costs on the labor supply decision, Zax and Kain (1991) showed that increases in commuting times generally increased the probability of employee quits.

Aside from the fixed costs of employment, employer inflexibility could be the factor that limits the days and hours that people are able to work, despite their preferences. Altonji and Paxson (1992) showed that married women who changed jobs exhibited more of a change in weekly and yearly hours compared with those who did not change jobs. They attributed this to employers restricting hours choices, which necessitated job change to attain the desired number of hours. Rettenmaier (1996) discovered that individuals who prefer low or high hours of work were more likely to be self-employed because they had a lower probability of finding these hours in paid jobs. In a related paper, Kahn and Lang (1995) found that over half of Canadians in 1985 were dissatisfied with the number of weekly hours they usually worked. Of these, about two-thirds expressed the desire for more weekly hours, not fewer. Of course, an increase in weekly hours can come from increasing days per week or hours per day or both. If fixed costs per day of work are high relative to hourly costs, we would find that these workers desired to put in more hours per day in increasing their hours per week. If the hourly costs of work are higher, we would expect the opposite, assuming that their are no employer-imposed constraints on the availability of hours and days.

There is, in fact, some evidence suggesting that the aggregation of days and hours into weekly hours results may result in poor labor supply estimates. Hamermesh (1996) provided estimates of the reduced-form correlates of days and daily hours in the absence of a formal model. He concluded that we cannot treat weekly hours as a reliable unit of analysis because daily hours and days per week both vary. He found that daily hours, in both the United States and Germany, tended to vary more than days per week in response to various exogenous shocks such as changes in the unemployment rate. This implies that the cost of changing days per week is higher than the cost of changing hours per day.

We want to dissect the weekly hours decision faced by workers. The first step will be to model this days and hours decision. If there are significant fixed costs to the number of hours per week and the number of weeks per year worked, fixed costs in daily and hourly terms may also be important in determining the combination of hours per day and days per week. The daily act of preparing for work and commuting to and from the work site results in substantial sunk costs that are borne by workers. In other words, are the hours per day and days per week decision joint? Are the determinants of the two the same? Are there significant costs per day or per hour of work which prevent people from seeking jobs? Or is it employers who constrain the available set of hours and days that employees may choose?

DATA

The 1990 Labour Market Activity Survey (LMAS) of 1990 will be utilized in the empirical part of the chapter. The LMAS is a unique data set that includes variables for days per week and hours per day usually worked—variables which are not normally found in labor force surveys.[3] Data on up to five jobs held by each individual in 1990 are also included. This will ultimately allow us to make inferences about the motives behind job change. The data set also includes other variables for the reason the respondent left the job and the number of additional monthly hours the respondent desired to work. This information will be useful in deducing whether it is fixed costs that result in various hours/days combinations, or whether it is rigidities in the labor market that do not make the desired combinations of hours/days available to employees.

The sample includes those between the ages of 17 and 64 who lived throughout the country, with the exception of the Northwest and Yukon territories. Those who did not hold any job in 1990 were eliminated from the sample, as were those who did not work at a paid job (i.e., the self-employed) or who attended school full-time at any time during the year.[4] Those who held more than two jobs in 1990 were eliminated. To avoid job overlap (due to moonlighting, for example), those who started a second job in a week preceding the completion of the first job were dropped, as were respondents who claimed to work more than 18 hours per day or to have earned less than $1.00 per hour at either job. Satisfying these criteria were 16,820 males and 14,635 females.[5] The sample is further disaggregated into job-stayers (14,577 males and 13,245 females) and job-changers who moved from one paid job to another paid job (1,563 males and 1,318 females).[6]

WEEKLY WORK PATTERNS

An initial look at the data reveals that hours of work tend to be more flexible than days of work. Table 1 gives the joint distribution of hours and days for males and females. We define the standard workday

Table 1 Joint Distribution of Hours per Day and Days per Week, All Workers, Paid and Self-Employed Workers, Males and Females (%)

Category/ hours per day Days per week	1–4	5	6–7	Total
Males				
All workers (*n* = 18,328)				
< 4.0	0.18	0.23	0.09	0.50
4.0–5.9	0.51	00.74	0.20	1.45
6.0–7.4	0.69	4.44	0.62	5.75
7.5–8.5	1.79	58.55	3.82	64.16
8.6–9.9	0.32	5.36	1.53	7.21
> 9.9	4.06	8.56	8.32	20.94
Total	7.55	77.88	14.58	100.00
Paid workers (*n* = 16,280)				
< 4.0	0.19	0.18	0.07	0.44
4.0–5.9	0.51	0.76	0.15	1.42
6.0–7.4	0.73	4.70	0.53	5.96
7.5–8.5	1.83	63.36	2.99	68.18
8.6–9.9	0.32	5.29	0.88	6.49
> 9.9	4.44	7.75	5.33	17.52
Total	8.02	82.04	9.95	100.00
Self-employed (*n* = 2,048)				
< 4.0	0.15	0.63	0.24	1.02
4.0–5.9	0.54	0.59	0.54	1.67
6.0–7.4	0.34	2.34	1.32	4.00
7.5–8.5	1.56	20.31	10.45	32.32
8.6–9.9	0.24	5.96	6.64	12.84
> 9.9	1.07	14.94	32.13	48.14
Total	3.90	44.77	51.32	100.00

(continued)

Table 1 (continued)

Category/ hours per day \ Days per week	1–4	5	6–7	Total
Females				
All workers (*n* = 15,263)				
< 4.0	1.13	1.47	0.33	2.93
4.0–5.9	3.97	4.48	0.56	9.01
6.0–7.4	5.37	15.20	0.89	21.46
7.5–8.5	8.25	46.28	2.44	56.97
8.6–9.9	0.39	2.10	0.55	3.04
> 9.9	2.37	2.14	2.06	6.57
Total	21.48	71.67	6.83	100.00
Paid workers (*n* = 14,635)				
< 4.0	1.11	1.42	0.23	2.76
4.0–5.9	4.00	4.48	0.50	8.98
6.0–7.4	5.46	15.54	0.77	21.77
7.5–8.5	8.41	47.47	2.15	58.03
8.6–9.9	0.38	2.00	0.42	2.80
> 9.9	2.44	1.98	1.23	5.65
Total	21.80	72.89	5.30	100.00
Self-employed (*n* = 628)				
< 4.0	1.75	2.55	2.71	7.01
4.0–5.9	3.34	4.46	1.91	9.71
6.0–7.4	3.35	7.32	3.66	14.33
7.5–8.5	4.29	18.63	9.39	32.31
8.6–9.9	0.64	4.46	3.50	8.60
> 9.9	0.80	5.89	21.34	28.03
Total	14.17	43.31	42.51	100.00

NOTE: Totals may not add due to rounding error.

to be in the 7.5 to 8.5 hour range and the standard workweek to be five days. For all male workers, 78 percent normally worked the standard five-day workweek. Only 64 percent of male workers worked the standard workday. As expected, female workers show greater diversity in their usual hours and days, as 72 percent worked the standard five-day week and 57 percent worked the "normal" workday.

Because we consider the self-employed as being somewhat less constrained by the days and hours restrictions of paid employees, we expect this group of workers to exhibit greater variance in their observed hours and days. Further breaking down the sample into paid workers and self-employed workers does in fact reveal this; i.e., paid workers tended to work more standard days and hours compared to the self-employed. Some 82 percent of paid males worked a five-day week in 1990, compared to only 45 percent of self-employed males. In fact, over 51 percent of self-employed males worked six- or seven-day weeks. Usual work hours were also more standardized for paid workers, with 68 percent working normal hours. By contrast, only 32 percent of self-employed males worked between 7.5 and 8.5 hours per day, with 48 percent working 10 hours per day or more. Both the hours and days distributions are more heavily weighted at the top for self-employed males. One interesting result is that over 4 percent of paid workers worked at least 10 hours per day but less than five days per week. This suggests that some workers were able to work longer hours and fewer days within a standard-length workweek.

Women generally show more flexibility in their hours and days combinations compared to men. For paid women, 73 percent worked a standard five-day week and 58 percent worked the standard workday—about 10 percentage points lower than the equivalent values for males. Paid females were also much more likely to work shorter hours and days than their male counterparts and less likely to work longer days and hours. As with the case of males, self-employed females showed much more variation in their hours and days; more were likely to work larger numbers of hours and days compared to female paid workers. They were also more concentrated in the lower tail of the hours per day distribution. Compared to self-employed males, females were more likely to work both shorter hours and days.

The patterns for both genders are generally consistent with those obtained by Hamermesh (1996) in his comparison of U.S. and German labor supply.[7]

A SIMPLE MODEL OF LABOR SUPPLY

Loosely extending a standard labor supply model such as the one found in Blank (1988), we can model the days and hours labor supply decision and then estimate the model. Each individual is assumed to maximize his or her utility, which is a function of the level of consumption and the amount of leisure consumed. Formally, the model can be written as

$$(1) \quad \max_{C,D_l H_l} U(C, D_l, H_l)$$

subject to

$$C = Y - \alpha D_w + D_w H_w \mathrm{W}(1 - \mu)$$

$$D_l = D - D_w$$

$$\mathrm{H}_l = \mathrm{D}_w (H - H_w),$$

where C is weekly consumption and is simply the amount of exogenous income available (Y) plus the amount of labor income earned per week. The latter is the usual number of days worked per week (D_w) times the usual number of hours worked per day (H_w) times the usual hourly wage (W). Finally, we subtract the costs of employment for both days of work and hours of work. The simple act of preparing for and commuting to work involves costs which are borne daily, regardless of the amount of time spent on the job. Other costs, however, are a function of the amount of time per day spent on the job. Costs such as day care and parking, for example, may be on an hourly basis. We assume that α represents the fixed costs per day of work, and μ are the costs per hour of work.

Leisure is divided into days per week of leisure (D_l), which is the number of days per week (D) less the number of days per week worked, and hours of leisure (H_l), which is the weekly amount of leisure consumed on days worked and is simply the number of days per week worked times the number of hours on these days not at work ($H - H_w$). Because hours and days of leisure may be qualitatively different for individuals, they enter the utility functions separately.

We know that solving the above problem yields Marshallian demand functions for hours of leisure and days of leisure, which can be transposed into labor supply functions for days per week of work and hours per day of work. In other words, we can solve for

(2a) $D_w = D_w(Y, W, \delta, \alpha, \mu)$

(2b) $H_w = H_w(Y, W, \delta, \alpha, \mu)$

where δ is a vector of demographic and job-related variables that we assume will affect supply for hours and days of work.

If we assume that leisure (in either days or hours) is a normal good, and that the substitution effect is greater than the income effect, then an increase in the cost of days or hours should increase the amount of leisure taken. Obversely, the number of hours and days of work supplied should decrease as the direct costs of each increase. Thus, we assume that $\partial D_w / \partial \alpha < 0$ and $\partial H_w / \partial \mu < 0$. Furthermore, if we assume that hours and days are substitutes, the cross-partial derivatives will both be positive. In other words, $\partial D_w / \partial \mu > 0$ and $\partial H_w / \partial \alpha > 0$ says that as the fixed cost per hour (day) of work increases, the individual will increase his or her supply of days (hours) because the opportunity cost of doing so is now relatively less expensive.

To operationalize the model into days and hours, we assume a linear approximation of the relationship between the supply of labor and its determinants. Thus, model (1) becomes

(3a) $D_w^* = X_1\beta_1 + \varepsilon_1$

(3b) $H_w^* = X_2\beta_2 + \varepsilon_2$

where $\boldsymbol{X}_1$ and $\boldsymbol{X}_2$ are vectors of individual and job characteristics that determine the number of days and hours supplied, β_1 and β_2 are the vector of coefficients, and ε_1 and ε_2 are the usual white noise error terms. Of course, D_w^* and H_w^* are only observed if the respondent is actually a labor force participant; they are written in natural logarithms.

Eqs. 3a and 3b, however, are limited because they implicitly assume that the hours and days decisions are separable. They also implicitly assume that there are no discontinuities in labor supply choices. It is well-established that discontinuities do in fact arise from the fixed costs of work (hourly, daily, weekly, etc.) as well as employer-imposed constraints which may limit the maximum or minimum hours that a person is able to work, thus narrowing the choice set of the worker. Still, it provides a starting point to analyze the determinants of hours and days of work. From estimation of Eqs. 3a and 3b, certain implications about daily and hourly costs of employment can be ascertained.

The reduced-from estimates of Eqs. 3a and 3b, with and without job controls, are presented in Table 2.[8] For economy of space, only the coefficients discussed are included (summary statistics and full results appear in Tables A1 and A2 in the appendix). At the bottom of the table, the results from Breusch-Pagan tests allow us to reject in all four cases the hypothesis that the days and hours regressions are independent. The correlation coefficients of the residuals are positive, underlining something we discovered in Table 1: workers who work, for unexplained reasons, more (fewer) hours also tend to work more (fewer) days. Still, the magnitudes of these correlation coefficients are small. Since the dependent variables are natural logarithms, we can compare the effect of the independent variables on hours and days. Significance levels of pairwise *t*-statistics, which test the hypothesis that the effect of the independent variable is the same on both hours and days, are also included. The coefficients on the number of children in the days and hours regressions, for example, are significantly different in the case of females but not in the case of males.

The results show that both men and women above 19 years of age work more hours and days compared to the control group of individuals between 17 and 19 years of age. For males, hours and days peak at 25–34 years of age, while hours and days for females reach a maximum at 20–24 years of age. Throughout the remainder of the life-

Table 2 OLS Estimates of ln (Hours) and ln (Days), With and Without Job Controls, Males and Females[a,b]

	Males						Females					
	No job controls			Job controls			No job controls			Job controls		
Independent variable	Hours	Days	t	Hours	Days	t	Hours	Days	t	Hours	Days	t
Age												
20–24 years	0.0493	0.0385		0.0501	0.0413		0.0327	0.0853	***	0.0248	0.0787	***
	(4.067)	(3.667)		(4.239)	(3.970)		(1.441)	(3.373)		(1.115)	(3.160)	
25–34 years	0.0549	0.0412		0.0537	0.0460		0.0225	0.0785	***	0.0010	0.0553	***
	(4.516)	(3.918)		(4.522)	(4.392)		(1.008)	(3.156)		(4.500)	(2.251)	
35–44 years	0.0371	0.0322		0.0427	0.0392		0.0106	0.0688	***	–0.0175	0.0385	***
	(3.006)	(3.007)		(3.518)	(3.667)		(0.467)	(2.735)		(0.787)	(1.542)	
45–54 years	0.0227	0.0310		0.0317	0.0377		–0.0176	0.0395	***	–0.0413	0.0118	***
	(1.792)	(2.832)		(2.530)	(3.414)		(0.763)	(1.541)		(1.816)	(0.463)	
55–64 years	–0.0165	0.0198	**	–0.0036	0.0269	*	–0.0807	–0.0362		–0.1025	–0.0637	
	(1.251)	(1.732)		(0.278)	(2.337)		(3.371)	(1.360)		(4.319)	(2.399)	
Number of children												
Ages 0–2	–0.0023	–0.0031		–0.0030	–0.0035		–0.0195	–.00742	*	–0.0163	–0.0681	*
	(0.579)	(0.893)		(0.752)	(1.003)		(2.867)	(9.788)		(2.430)	(9.102)	
Ages 3–5	0.0059	0.0005		0.0060	0.0012		–0.0332	–0.0645	*	–0.0320	–0.0616	*
	(1.507)	(0.152)		(1.590)	(0.354)		(5.153)	(9.012)		(5.078)	(8.726)	

(continued)

Table 2 (continued)

	Males						Females					
	No job controls			Job controls			No job controls			Job controls		
Independent variable	Hours	Days	*t*	Hours	Days	*t*	Hours	Days	*t*	Hours	Days	*t*
Ages > 5	0.0022	0.0007		0.0011	0.0005		–0.0192	–0.0277	*	–0.0167	–0.0224	***
	(1.445)	(0.559)		(7.250)	(0.348)		(7.749)	(10.082)		6.843	8.206	
Constant	2.1163	1.5883	*	2.1341	1.5712	*	1.9751	1.5201	*	1.9535	1.4841	*
	(151.034)	(130.867)		(127.509)	(106.577)		(75.150)	(52.004)		61.491	(41.755)	
R^2	0.0313	0.0086		0.0841	0.0305		0.0276	0.0696		0.0408	0.0713	
Correlation coefficient of residuals	0.1455			0.1379			0.1545			0.1371		
Breusch–Pagan test of independence (p-value)	0.0000			0.0000			0.0000			0.0000		
Number of observations	16,280			16,280			14,635			14,635		

[a] Absolute values of *t*-statistics are in parentheses.

*** = Significant at 1% level.

** = Significant at 5% level.

* = Significant at 10% level.

[b] Controls for marital status, relationship to family head, education, region, mother tongue, immigrant, and visible minority were all included in all regressions. Job control variables are firm size, industry, occupation, union coverage, tenure, and pension coverage. Age 17–19 is the omitted variable.

cycle, labor supply in both dimensions declines slowly. For both genders, the increase in weekly hours is the result of a larger percentage increase in days than hours, but only in the case of women are these differences statistically significant at 10 percent.

These results are somewhat at odds with those of Hamermesh (1996), who found that the inverse U-shaped pattern was steeper in the case of hours than days for both U.S. and German male and female workers. He reasoned that the steeper hours profile implied that hours are less costly to add than days. In our case, male workers do add hours slightly more rapidly than days until they reach their peaks between 25 and 34 years. Thereafter, hours fall off more rapidly than days. Because 82 percent of the males in our sample work a standard five-day workweek, flexibility in weekly hours obviously comes from changes in hours. For female workers, it is days that rise more rapidly to a peak at 20–24 years, and then both days and hours decline at a similar rate. Thus females appear to be more flexible in altering both days and hours than males.

Because women still generally hold the primary responsibility for child care, the presence of young children should decrease their labor supply. Indeed, the number of young children present does have a negative effect on supply of both hours and days. In the case of no job controls, the point estimates show that the presence of a child two years of age or less is related to a drop in hours of almost 2 percent but a drop in days of over 7 percent. By contrast, the presence of young children has no statistically significant effect on the male supply of hours and days. Thus, the data show that young children are correlated with a decline both the hours and days supplied by the mother. If the fixed costs of child care are incurred on a daily basis versus an hourly basis, we would expect the decline in labor supply to be borne by a larger decline in days. This indeed is the case.

JOB-CHANGE BEHAVIOR AND HOURS PER DAY AND DAYS PER WEEK

Insofar as hours and days combinations within jobs are less than optimal for workers, we might expect job change to occur in order to attain the desired combination. Since the LMAS contains data on up to five job changes per worker per year, we can use this information to estimate a model of job change behavior that will help give further insights into the daily and hourly fixed costs of labor supply.

Modifying the labor force participation model of Blank (1988), we assume that an individual i in period j attains the utility level $U_i^j(H_i^j, D_i^j, E_i^j)$, where H_i^j is hours of work per day at the job held in period j, D_i^j is days per week at job j, and $E_i^j = H_i^j \cdot D_i^j \cdot W_i^j$ represent the weekly earnings at job j, which has an hourly wage rate of W_i^j. If we assume that the individual has perfect information about all the arguments in his or her utility function, and that job mobility is cost-free, then an individual will change jobs only if

$$\textbf{(4)}\quad P^* = U_i^2(H_i^2, D_i^2, E_i^2) - U_i^1(U_i^1, D_i^1, E_i^1) \geq 0.$$

If we further assume that P^* is linearly dependent on the three arguments in the utility function, as well as other demographic and economic variables that determine job change, we can write

$$\textbf{(5)}\quad P^* = \lambda \mathbf{Z} + v,$$

where $\mathbf{Z}$ is the aforementioned vector of job change determinants, λ is its corresponding vector of coefficients, and v is the white noise error term. Utility is unobservable and therefore so is the variable P^*. What we do observe, however, is a dichotomous variable P, where

$$P = 1 \text{ if } P^* \geq 0 \text{ and } P = 0 \text{ if } P^* < 0.$$

If we assume that an individual is in equilibrium at the initial job, utility is being maximized. A shock which affects one or more arguments in the individual's utility function may result in the utility no longer being maximized at that job, and job change will occur if utility can be maximized at a new job. The birth of a child, for example increases the fixed costs of employment along both time dimensions. If the fixed costs of hours are more costly than the fixed costs of days, we would expect job-changers to want to move out of jobs with longer hours. Conversely, if days are more costly than hours, we would expect the probability of leaving to be positively related to the days variable.

Probit estimates of Eq. 5 appear in Table 3. Again, only the relevant coefficients are included with full results contained in Table A3 in the appendix. Independent variables include all personal and job controls used in the previous analysis, plus variables for ln (hours) and ln (days) at the initial job, a dummy which equals 1.0 if respondents said that they desired more monthly hours at their initial job, and interactions of this variable with ln (hours) and ln (days). Separate probits are also estimated without personal and job controls. Of the 16,280 males in our sample, some 1,563 changed jobs in 1990 for a probability of 0.096. For women, the probability of moving from a first to a second job was 0.090 (1,318 changers from a sample of 14,635).

The effects of personal and job characteristics on job change are very similar for both genders. Probability of job change decreases as age increases. The presence of small children also reduces the likelihood of job change for both genders, especially women. Older children have little influence on the job change behavior of men but continue to slightly lessen female mobility. Union members tend to exhibit less job mobility as do those with more job tenure. These results are consistent with the literature.[9]

The marginal effects of hours, days, and wanting extra monthly hours on job-change probability are presented in Table 4. For males, the number of hours and days worked at their initial jobs have little effect on job-change probability when control variables are included. In the case of no controls, the coefficient on ln (hours) is twice as large as that on ln (days), although only the former is significant. The effect of the desire for extra hours, however, is significant in both cases, accounting for about half of the predicted probability when controls are not included. For females, a different pattern emerges as the ln (days) and ln (hours) coefficients are similar in magnitude and significant in each instance. The effect of the desire to work extra hours is significant in each case and large compared to the predicted probabilities of each model.

These results are not supportive of the hypothesis that the addition of days is more costly than the addition of hours, but that could simply be a result of unobserved heterogeneity between workers, which causes them to separate from their initial jobs if their preferences do not favor the given hours/days combination. Regardless of the reason, this result shows that the desire to work extra hours significantly increases the

Table 3 Probit Estimates of Job Change Probability, Males and Females[a,b]

	Males				Females			
	No controls		With controls		No controls		With controls	
Independent variable	Coefficient	Partial deriv.	Coefficient	Partial deriv.	Coefficient	Partial deriv.	Coefficient	Partial deriv.
ln (hours per day)	0.2869 (3.442)	0.048	0.0806 (0.890)	0.010	0.3203 (4.394)	0.051	0.2358 (2.997)	0.027
ln (days per week)	0.1946 (1.832)	0.033	–0.0321 (0.296)	–0.004	0.3093 (4.522)	0.049	0.2330 (3.188)	0.027
Extra hours wanted	1.7816 (4.231)	0.548	0.5529 (1.263)	0.092	0.8750 (2.913)	0.208	0.5999 (1.919)	0.098
Extra hours wanted × ln (hours)	–0.2810 (1.515)	–0.047	0.0388 (0.199)	0.005	–0.1086 (0.791)	–0.017	–0.0309 (0.213)	–0.004
Extra hours wanted × ln (days)	–0.5910 (3.518)	–0.100	–0.3389 (1.950)	–0.041	–0.1856 (1.651)	–0.029	–0.2079 (1.763)	–0.024
Age								
20–24 years			–0.3188 (3.487)	–0.031			–0.0973 (0.836)	–0.011

25–34 years	−0.4625	−0.049	−0.2862	−0.031
	(4.925)		(2.459)	
35–44 years	−0.5028	−0.052	−0.4378	−0.045
	(5.145)		(3.641)	
45–54 years	−0.6083	−0.055	−0.5785	−0.051
	(5.852)		(4.616)	
55–64 years	−0.8859	−0.063	−0.7840	−0.055
	(7.612)		(5.477)	
Number of children				
Ages 0–2	−0.1106	−0.013	−0.2743	−0.032
	(2.771)		(5.783)	
Ages 3–5	−0.0032	0.000	−0.0619	−0.007
	(0.085)		(1.425)	
Ages > 5	−0.0196	−0.002	−0.0345	−0.004
	(1.242)		(2.015)	
Covered by union agreement	−0.2124	−0.025	−0.2854	−0.032
	(5.768)		(6.611)	
Tenure at job (weeks/100)	−0.1179	−0.014	−0.1460	−0.017
	(18.130)		(15.541)	

(continued)

Table 3 (continued)

	Males				Females			
	No controls		With controls		No controls		With controls	
Independent variable	Coefficient	Partial deriv.	Coefficient	Partial deriv.	Coefficient	Partial deriv.	Coefficient	Partial deriv.
Constant	–2.2614 (9.773)		–0.3095 (1.096)		–2.5003 (14.577)		–1.6369 (6.028)	
Covered by union agreement								
Pseudo R^2	0.007		0.124		0.012		0.119	
X^2(5, 55, 5 and 54 d.f.)	69.32		1277.03		110.18		1055.43	
Observed P		0.096		0.096		0.090		0.090
Predicted P		0.095		0.061		0.087		0.058
Number of observations	16,280		16,280		14,635		14,629	

[a] Absolute values of *t*-ratios are in parentheses.

[b] Controls are marital status, relationship to family head, education, region, mother tongue, immigrant, visible minority, firm size, industry, occupation, and pension coverage. Age 17–19 is the omitted variable.

Table 4 **Effect of Independent Variables on Predicted Job Change Probability**[a]

	Males		Females	
	w/o controls	w/controls	w/o controls	w/controls
Hours	0.052 (3.372)	0.010 (0.932)	0.060 (4.672)	0.032 (3.257)
Days	0.026 (1.415)	–0.008 (0.661)	0.055 (4.691)	0.029 (3.178)
Extra hours wanted	0.048 (5.388)	0.011 (1.921)	0.076 (8.222)	0.032 (4.580)
Predicted *P*	0.095	0.061	0.087	0.058

[a] Absolute values of *t*-ratios are in parentheses.

probability of job change for both genders, suggesting that those who want these hours may face hours and/or days constraints in their initial jobs and have to change jobs to relax these constraints.

The fact that extra hours are desired indicates that the individual may not be in equilibrium at his or her current job. The interactions of the extra hours variable with the days and hours variables in Table 3 provide us with additional information about the reasons for job change. If a lack of hours (or days) on the current job is the constraint on lower-than-desired total hours and thus at least part of the motivation for seeking a new job, then we would expect a negative coefficient on the interaction term. For example, if one wanted extra days per week, the probability of job change should decrease as the number of days at the current job increase. The significant negative coefficients on the extra hours/days interaction variables in the case of males support this assertion. For the extra hours/hours interaction, however, we cannot reject the null of no effect on job-change behavior. Thus, a male who wants extra hours is less likely to leave his current job as days per week increase, while hours per day have no significant effect on job change. For females, none of the coefficients on either of the interaction variables is statistically significant.

These results suggest that for males who do not desire extra hours, the costs of additional hours of work are higher, at least within their first jobs, than those of additional days of work because the increase in job-change probability is greater for an increase in hours relative to a comparable increase in days. For females, increases in both days and hours result in similar changes in job-change probabilities. For both who want more monthly hours, a larger number of hours at the worker's initial job has a smaller negative effect on job-change behavior than an equivalent increase in days. Thus, the probability of job change decreases as extra hours and days are added within the initial job, although the negative effect of the latter is larger. This appears to be inconsistent with the hypothesis of higher daily fixed costs, although it may simply imply that the monetary benefits to working extra days far outweigh the fixed costs of these days. It may also mean that job-changers are somewhat constrained in their choices, as new job offers are more apt to include a higher number of days rather than a higher number of hours.[10]

These estimates do give us insight into what motivates job change, but they tell us little about the changes in hours and days that result from job change. This is addressed in the following section.

DETAILED ANALYSIS OF JOB-CHANGERS

Tables 5 and 6 show the extent to which the means and variances of ln (days) and ln (hours) vary as workers change jobs. In both these tables, columns 1 through 4 display the means and variances of ln (hours) and ln (days) for each of the two jobs held by job-changers, and the one job held by job-stayers. Column 5 shows the mean changes in log weekly hours as well as the corresponding variances. Columns 6 and 7 disaggregate weekly hours changes into changes in ln (hours) and ln (days). Columns 8 and 9 display the test statistics for differences in the means and variances in ln (hours) and ln (days) changes.[11] We also disaggregate the sample of job-changers, first into voluntary and involuntary job-changers, and then into those who wanted extra hours at their first job and those who did not.[12] In each of these two subsamples, significant differences exist between the two groups in terms of mean changes and variances in almost every time dimension. For example, both the means and variances of changes in weekly hours are significantly different when comparing those who desired extra

Table 5 Selected Means and Variances of ln (hours) and ln (days), Jobs 1 and 2 and First Differences, for Male Job-Stayers and Job-Changers[a]

								Tests for differences in[b]	
	Job 1		Job 2		First differences			Means	Variance
	ln (hours)	ln (days)	ln (hours)	ln (days)	Δln (wkly. hours)	Δln (hours)	Δln (days)	z	F
Job stayers	2.116	1.602	N/A	N/A	N/A	N/A	N/A		
(n = 14,577)	(0.0341	(0.0234)	(N/A)	(N/A)	(N/A)	(N/A)	(N/A)		
Job changers	2.128	1.596	2.132	1.598	0.0067	0.0040	0.0027	0.146	
(n = 1,563)	(0.0395)	(0.0423)	(0.0433)	(0.0394)	(0.1497)	(0.0578)	(0.6640)		1.149**
Voluntary	2.119	1.589	2.125	1.603	0.0190	0.0058	0.0131	0.709	
(n = 1,000)	(0.0429)	(0.0389)	(0.0412)	(0.0276)	(0.3548)	(0.0548)	(0.0511)		1.072
Involuntary	2.142	1.607	2.143	1.591	–0.0150	0.0007	–0.0158	0.990	
(n = 563)	(0.0331)	(0.4830)	(0.0470)	(0.0602)	(0.1916)	(0.0632)	(0.0933)		1.476*
Extra hours wanted	2.045	1.459	2.117	1.570	0.1824	0.0714	0.1110	1.400	
(n = 230)	(0.0584)	(0.0942)	(0.0379)	(0.0390)	(0.2134)	(0.0686)	(0.1155)		1.684*
Extra hours not wanted	2.142	1.619	2.134	1.603	–0.0236	–0.0076	–0.0160	0.921	
(n = 1,333)	(0.0349)	(0.0299)	(0.0442)	(0.0393)	(0.1326)	(0.0551)	(0.0557)		1.011

[a] The top data row of each pair is the mean (μ) and the second row, in parentheses, is the variance (σ^2).

[b] * and ** denote significance at the 1% and 5% levels, respectively.

Table 6 Selected Means and Variances of ln(Hours) and ln(Days), Jobs 1 and 2 and First Differences, for Female Job-Stayers and Job-Changers[a]

	Job 1		Job 2		First differences			Tests for differences in[b]	
								Means	Variance
	ln (hours)	ln (days)	ln (hours)	ln (days)	Δln (wkly. hours)	Δln(hours)	Δln(days)	z	F
Job stayers	1.967	1.491	N/A	N/A	N/A	N/A	N/A		
(n = 13,245)	(0.0813)	(0.0100)	(N/A)	(N/A)	(N/A)	(N/A)	(N/A)		
Job changers	1.996	1.520	1.997	1.507	–0.0113	0.0008	–0.0121	1.624	
(n = 1,318)	(0.0469)	(0.0784)	(0.0547)	(0.0909)	(0.2429)	(0.0732)	(0.1287)		1.758*
Voluntary	1.991	1.509	1.998	1.518	0.0162	0.0071	0.0091	0.146	
(n = 998)	(0.0462)	(0.0814)	(0.0453)	(0.0787)	(0.2306)	(0.0690)	(0.1179)		1.709*
Involuntary	2.012	1.553	1.993	1.475	–0.0969	–0.0188	–0.0782	2.157**	
(n = 320)	(0.0488)	(0.0670)	(0.0841)	(0.1280)	(0.2724)	(0.0858)	(0.1569)		1.829*
Extra hours	1.862	1.329	1.985	1.483	0.2768	0.1225	0.1542	0.845	
wanted	(0.0814)	(0.1692)	(0.0635)	(0.0989)	(0.3967)	(0.1072)	(0.2275)		2.122*
(n = 238)									
Extra hours	2.026	1.562	2.000	1.513	–0.0747	–0.0260	–0.0487	2.299**	
preferred	(0.0345)	(0.0487)	(0.0528)	(0.0891)	(0.1870)	(0.0618)	(0.0996)		1.612*
(n = 1,080)									

[a] The top data row of each pair is the mean (μ) and the second row, in parentheses, is the variance (σ^2).

[b] * and ** denote significance at the 1% and 5% levels, respectively.

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monthly hours with those who did not.[13] The evidence appears to support the hypothesis that flexibility in weekly work schedules is important to employees.

Males who changed jobs worked slightly more hours per day and marginally fewer days per week at their initial jobs compared with those who did not change jobs, although only the former difference is statistically significant. The variances of both ln (hours) and ln (days), however, are significantly higher in the case of job-changers than job-stayers. Thus, although we cannot reject the hypothesis that the average job-changer works the same number of daily hours as the average job-stayer, job-changers show much more variation in their work schedules at their initial jobs, and this variation persists as they move into new jobs. This is generally consistent with the results of Altonji and Paxson (1986), who found that the variance of time worked in a number of dimensions (hours/week, weeks/year, or hours/year) increased for people who changed jobs, relative to those who did not change jobs.[14]

For all male job-changers, column 5 of Table 5 shows that job-changers worked marginally more weekly hours at their new jobs. Disaggregating these changes into changes in hours and days, however, we cannot reject the hypothesis that the change in weekly hours was the result of equal changes in both hours and days. The variance of days changes, however, is higher than that of hours at the 5 percent level. Voluntary changers had an increase in ln (weekly hours) of 0.019, with the adjustment coming equally from increases in hours and days. Involuntary changers, by contrast, had a decline of 0.015 in ln (weekly hours). Although we cannot reject the hypothesis that the change came equally from changes in hours and days, we see that the variance of the change in days is significantly higher than the change in hours.

Males who wanted extra hours had an increase in ln (weekly hours) of 0.182 compared to a decline of 0.024 for those who did not. Only in the former case, however, is the variance of the change in ln (days) significantly higher than that of the change in ln (hours). The higher variances for those desiring extra hours also imply a great deal of flexibility in finding the preferred combinations of hours and days. Similar results are obtained by Altonji and Paxson (1988, 1992), who found that workers who desired to work more weekly hours were more likely to increase these hours when they changed jobs.

The results for females are presented in Table 6 and show much more variation in all time dimensions compared to males. Comparing the distributions of hours and days on the first job between stayers and changers, both means are significantly larger for job-changers. For job-changers, the variance in log hours is significantly smaller than for stayers while the opposite holds for the variance on ln (days). Female job-changers find new jobs with only marginally more hours and marginally fewer days compared to their previous positions, in neither case are the differences significant. Changers do, however, move into jobs with significantly higher variances in both days and hours.

For all female job-changers, average total weekly hours declined slightly following job change, although the changes in hours and days are statistically indistinguishable. The variance of hours changes, however, is statistically much smaller than the variance of days changes. Voluntary changers had a modest increase in weekly hours compared to involuntary changes. Only in the latter case, however, can we say that the bulk of this change was the result of the steep decline in days. Those women who wanted extra hours increased their ln (weekly hours) by an average of 0.277, the result of an equal increase in both hours and days. Those who did not want extra hours saw their mean ln (weekly hours) decline by 0.075, with most of this decline the result of a drop in days. In each case, the variance of the days change is statistically larger than that of the hours change.

Both male and female job-changers experience changes in days and hours as they change jobs. Mean differences, however, appear to hide important details. In many cases we cannot say with certainty that the adjustment in weekly hours is the result of either changes in days or hours. The fact that the variance of changes in days is frequently significantly larger in the case of males, and is always significantly larger in the case of females, shows that the changes in days are much more flexible than changes in hours when workers change jobs. This then implies that changes in days are less flexible than changes in hours within jobs. These results are generally inconsistent with those obtained by Hamermesh (1996) for German and American workers. Only in the case of German males did the variance of the change in ln (days) exceed that of the change in ln (hours). For German women and American men the opposite held, while there was little difference for American women.[15]

In sum, we have seen that job-changers have a larger number of hours and days, generally with higher variances at their initial jobs compared with those who do not change jobs. Job change does not increase the average number of hours or days amongst changers but, with the exception of changes in male days, variances also increase. Voluntary job-changers of both genders experience higher average increases in both days and hours than those who do not change jobs voluntarily, although the variances in the changes of log hours and log days are larger for involuntary movers. Finally, those desiring extra hours at their initial jobs are likely to find these extra hours in the form of both more daily hours and more days per week, with the variances here (i.e., on the new job) smaller for those who did not want extra hours.

In terms of the daily and hourly costs of employment, the above provides weak evidence that the latter may be proportionately less than the former. Only in two cases (female involuntary changers and females not desiring extra hours) are the differences between mean ln (hours) and mean ln (days) changes statistically significant. In both of these cases, the percentage decline in hours is much smaller than the percentage decline in days. This suggests that hours of work are less costly to eliminate than days of work as these women changed jobs. Conversely, if we make the reasonable assumption of symmetry of costs in adding and subtracting days or hours, then adding extra days are more costly than adding extra hours.

The above results give us a good idea of the differences in mean hours and days changes along with their dispersions as job-changers move from one job to the next. This analysis, however, depends on differences in average hours and days changes between groups and thus could cloud the direction of changes in hours and days as workers move between jobs. Looking at the direction and magnitude of changes in hours and days by job-changers may offer some insights into what motivates job change. If daily fixed costs are indeed relatively high compared to hourly costs, we would expect changers to sort into new jobs with fewer days and more hours, all other things equal. Furthermore, we might expect these changes to be especially pronounced in the case of voluntary job-changers because they presumably are less constrained in their choices than those who change involuntarily. In other words, the voluntarily displaced worker may only change jobs if the hours/days combination at the new job is sufficiently different from the original job, a choice which an involuntarily displaced worker may not have. We might also expect those who want extra hours to attain these hours by increasing their daily hours rather than by increasing days. Females, who have exhibited more flexibility in our sample, might also be more apt to change into jobs with relatively fewer days and more hours.

Tables 7 and 8 show contingency tables for days and hours changes for male and female job-changers. In all but one case, χ^2 values allow us to reject the hypothesis that the distributions of days and hours are independent.[16] In other words, changes in hours and days between jobs are not purely random. In each of the 10 panels in Tables 7 and 8, no change in days is more likely than no change in hours for job-changers. Thus, it appears that rigidities in days are more prevalent than changes in hours. If larger fixed costs are incurred on a daily basis, these results are what we would expect.

The second and third panels of Table 7 show that there are differences between voluntary and involuntary job-changers. If voluntary job-changers are able to sort into a more palatable days/hours combination, and if the costs of adding a day of work are higher than adding more hours, we would expect positive hours changes to be more likely as workers move between jobs. If we simply look at aggregate days and hours changes, this does not appear to be the case. Although voluntary changers are more likely to hold on to both their original days and hours, the distributions show few other differences, if we only look at column and row totals. What is interesting is the off-diagonal elements of each panel. In this case voluntary changers were about as likely to increase days and decrease hours as they were to do the opposite. Involuntary changers, however, were more likely to increase hours and decrease days than to decrease hours jointly with increasing days. This is the opposite of what we expected, although somewhat supportive of the higher daily fixed costs hypothesis. This also suggests that factors other than the costs of days and hours exert more of an influence on voluntary job-changers as they sort into new jobs.

Those who said they wanted additional hours at their first jobs do, however, display important differences compared with those who did not want extra hours. Job-changers were more likely to have both positive hours and days changes if they wanted extra hours than if they did

Table 7 Changes in Hours and Days for Male Job-Changers (%)

	Δ Hours					
Δ Days	≥ 2.0	0.1– 2.0	0	< 0 & >–2	≤ –2	Total
All job-changers (n = 1563)[a]						
> 1	2.94	0.70	1.41	0.38	0.70	6.14
1	2.82	1.15	3.20	0.90	1.60	9.66
0	4.73	5.63	49.07	5.76	4.67	69.87
–1	2.30	0.45	2.11	1.15	1.98	8.00
< –1	1.15	0.13	2.05	0.38	2.62	6.33
Total	13.95	8.06	57.84	8.57	11.58	100.00
Voluntary job-changers (n = 1000)[b]						
> 1	3.30	0.40	1.10	0.50	0.80	6.10
1	2.40	1.40	3.10	1.10	1.30	9.30
0	4.20	5.70	51.00	6.40	4.60	71.90
–1	2.30	0.40	1.40	1.20	2.10	7.40
< –1	0.80	0.10	2.20	0.50	1.70	5.30
Total	13.00	8.00	58.80	9.70	10.50	100.00
Involuntary job-changers (n = 563)[c]						
> 1	2.31	1.24	1.95	0.18	0.53	6.22
1	3.55	0.71	3.37	0.53	2.13	10.30
0	5.68	5.51	45.65	4.62	4.80	66.25
–1	2.31	0.53	3.37	1.07	1.78	9.06
< –1	1.78	0.18	1.78	0.18	4.26	8.17
Total	15.63	8.17	56.13	6.57	13.50	100.00
Extra hours wanted (n = 230)[d]						
> 1	6.96	1.30	3.48	2.61	2.17	16.52
1	6.96	2.61	3.04	1.30	1.30	15.22
0	6.09	3.48	41.74	4.78	2.17	58.26
–1	3.04	0.00	0.87	0.87	1.30	6.09
< –1	0.87	0.00	2.17	0.00	0.87	3.91
Total	23.91	7.39	51.30	9.57	7.83	100.00

(continued)

Table 7 (continued)

Δ Days	Δ Hours ≥ 2.0	 0.1– 2.0	 0	 < 0 & >–2	 ≤ –2	Total
Extra hours not wanted (n = 1080)[e]						
> 1	2.25	0.60	1.05	0.00	0.45	4.35
1	2.10	0.90	3.23	0.83	1.65	8.70
0	4.50	6.00	50.34	5.93	5.10	71.87
–1	2.18	0.53	2.33	1.20	2.10	8.33
< –1	1.20	0.15	2.03	0.45	2.93	6.75
Total	12.23	8.18	58.96	8.40	12.23	100.00

[a] χ^2 (16 d.f.) = 400.33 (p = 0.000)
[b] χ^2 (16 d.f.) = 278.68 (p = 0.000)
[c] χ^2 (16 d.f.) = 151.88 (p = 0.000)
[d] χ^2 (16 d.f.) = 73.98 (p = 0.000)
[e] χ^2 (16 d.f.) = 336.48 (p = 0.000)

Table 8 Changes in Hours and Days for Female Job-Changers (%)

	Δ Hours					
Δ Days	≥ 2.0	0.1– 2.0	0	< 0 & >–2	≤ –2	Total
All job-changers (*n* = 1318)[a]						
> 1	2.58	1.44	2.35	0.53	1.37	8.27
1	2.43	1.29	2.35	0.61	1.52	8.19
0	4.86	8.27	39.45	4.48	8.35	65.40
–1	1.21	1.29	2.50	0.46	2.43	7.89
< –1	0.91	1.21	3.19	0.46	4.48	10.24
Total	11.99	13.51	49.85	6.53	18.13	100.00
Voluntary job-changers (*n* = 998)[b]						
> 1	2.71	1.70	2.51	0.60	1.40	8.92
1	2.40	1.00	2.20	0.50	1.60	7.72
0	4.61	8.12	42.28	5.31	8.12	68.44
–1	1.20	1.40	2.10	0.30	1.80	6.81
< –1	0.90	0.80	2.61	0.40	3.41	8.12
Total	11.82	13.03	51.70	7.11	16.33	100.00
Involuntary job-changers (*n* = 320)[c]						
> 1	2.19	0.62	1.88	0.31	1.25	6.25
1	2.50	2.19	2.81	0.94	1.25	9.69
0	5.62	8.75	30.63	1.88	9.06	55.94
–1	1.25	0.94	3.75	0.94	4.38	11.25
< –1	0.94	2.50	5.00	0.62	7.81	16.88
Total	12.50	15.00	44.06	4.69	23.75	100.00
Extra hours wanted (*n* = 238)[d]						
> 1	9.24	5.04	5.88	1.26	1.68	23.11
1	6.72	3.78	6.72	1.68	1.68	20.59
0	9.24	6.30	15.97	1.26	5.88	38.66
–1	2.10	2.10	3.36	0.84	2.10	10.50
< –1	2.52	0.84	0.84	0.42	2.52	7.14
Total	29.83	18.07	32.77	5.46	13.87	100.00

(continued)

Table 8 (continued)

Δ Days	Δ Hours					Total
	≥ 2.0	0.1– 2.0	0	< 0 & >–2	≤ –2	
Extra hours not wanted (*n* = 1080)[e]						
> 1	1.11	0.65	1.57	0.37	1.30	5.00
1	1.48	0.74	1.39	0.37	1.48	5.46
0	3.89	8.70	44.63	5.19	8.89	71.30
–1	1.02	1.11	2.31	0.37	2.50	7.31
< –1	0.56	1.30	3.70	0.46	4.91	10.93
Total	8.06	12.50	53.61	6.76	19.07	100.00

[a] χ^2 (16 d.f.) = 209.66 (p = 0.000)
[b] χ^2 (16 d.f.) = 164.58 (p = 0.000)
[c] χ^2 (16 d.f.) = 53.19 (p = 0.000)
[d] χ^2 (16 d.f.) = 21.42 (p = 0.163)
[e] χ^2 (16 d.f.) = 166.24 (p = 0.000)

not. Again, the off-diagonal elements show that these job changes had a higher propensity to accept more days in combination with fewer hours rather than fewer days and more hours. These results are contrary to our expectations.

Female job-changers show more flexibility in changing work schedules compared with males. This is reflected by the fact that they are less likely to move between jobs with identical hours and days pairs. Subsamples within female job-changers exhibit similar patterns to those of male job-changers. The off-diagonal elements show that in each of the five cases, female job-changers are marginally more likely to move into jobs with more hours and fewer days than into jobs with more days and fewer hours. This evidence is mildly supportive of our hypothesis of higher fixed costs of working more days.

In sum, these results suggest that there is a great deal of rigidity in job schedules. This is especially true of days per week since job-changers are less likely to change days than to change hours. As we have already discussed above (Table 1), females generally have more flexibility in their work schedules than men. Job change simply increases this flexibility.

CONCLUSIONS

By disaggregating weekly labor supply into hours per day and days per week, we have learned several interesting things. In bivariate distributions, employees tend to be clustered around standard hours and days, with men on average exhibiting less flexibility in hours and days worked than women. The self-employed are much less likely to work standard hours and days. Most of the self-employed males work both more hours and days compared to paid employees. Self-employed females also work more hours and days, but they are also more likely to work fewer hours per day.

The OLS estimates of hours and days supplied show that women with young children supply less labor in both dimensions, although the percentage drop in days is significantly larger. This implies that child care costs are borne on a daily basis, or at least are higher on a daily basis than an hourly basis.

Probit estimates show that a larger number of hours at the worker's initial job are significantly related to an increased probability of job change for both genders. Days per week are also a significant determinant of job change for females but have little effect on male job change behavior. For both genders, the desire for extra hours is related to increased probability of job change. As extra days are added at the initial job, this probability declines. The addition of hours at the initial job, however, has no significant effect on job change among those who want extra monthly hours. These results do not generally support the hypothesis that daily costs of employment are higher than hourly costs, although these costs may have been outweighed by the benefits of working more days, which may be the motivating factor behind job change.

A more detailed analysis of hours and days changes reveals that job-changers desire flexibility in their weekly work schedules. Not only do job-changers show more variability in their days and hours compared to job-stayers at their first jobs, but this carries over to their new jobs as well. The wider distribution in days changes compared to hours changes implies that flexibility in days is more important to job-changers than flexibility in hours. That working standard days is more common than working standard hours in the sample simply underlines the importance of flexibility in days, at least amongst job-changers.

The direction of hours and days changes suggests that there is a great deal of rigidity in weekly work schedules, especially in terms of days. Male job-changers are much more likely than females to move into new jobs with the same or very similar days and hours combinations. The fact that women are marginally more likely to increase hours and decrease days than to do the opposite, is mildly supportive of the hypothesis of higher daily fixed costs.

On the basis of the evidence presented above, we cannot say conclusively that the daily or hourly costs of employment drive the behavior of individuals, although they do appear to be influential. We can conclude that workers do desire more flexibility in their choice of hours and days. This is particularly true of days. Our analysis also points to the difficulty that workers may encounter in attaining optimal hours and days combinations. What we can conclude with more certainty is that using weekly hours, or more aggregated labor supply measures, hides important differences in the labor supply decisions of individuals.

Underlying the fact that the common labor supply aggregates hide important details are a host of policy implications. We have argued that the costs of employment are largely incurred on a daily rather than an hourly basis. Workers must get out of bed, go through the physical preparations for work, prepare lunches, get the kids ready for the day ahead and transport them to daycare or school before braving the daily commute to the worksite. This all happens before they actually do or are paid for work in the market. At the end of the day, this scenario is largely reversed. Every one of these costs is incurred on a daily basis, regardless of the number of hours actually worked. While little can be done about most of these daily costs, child care and commuting do have important policy implications for employers and for government.

The importance of young children in females' labor force participation decisions is already well-known. The results above also show that the presence of young children influences labor supply differently along the hours and days dimensions. This suggests that policies that reduce the daily costs of child care might be more important than those that reduce the hourly costs. Company provision of child care facilities at the worksite, for example, could result in significant savings to parents in terms of the time and dollar expenses of delivering children to an outside facility before commuting to work. For the employer, a better understanding of workers' time preferences could mean lower turnover rates, thus lowering associated costs. It could also mean less absenteeism and enhanced productivity from employees if they are able to work their desired hours/days combinations.

The increase in flexible working schedules has arguably been useful in reducing rush-hour traffic congestion in many North American cities (although increasing the length of the "hour"). This has likely resulted in reduced daily commute times and the costs of traveling to and from the worksite. Increased flexibility could further reduce these daily commuting costs. As cities continue to spread out over larger geographical areas, commute times, and the expenses associated with them, may also grow. A one-hour commute to the worksite in any major urban center in Canada is no longer considered unreasonable. The increased direct and indirect costs associated with longer commutes, along with the growth in appropriate technology, are undoubtedly reasons for the increased popularity of telecommuting. What does having the ability to work at home imply about the desired hours/days combination? Obviously the costs of both time dimensions decrease, but what is the optimal combination for the employee?

Related to this are public policy decisions regarding the provision of roadways and public transportation. Such decisions could benefit from a better understanding of days and hours preferences. If employees prefer to have more flexible daily hours and work fewer days per week, a large investment in public transportation facilities would not be warranted as the number of daily trips would be reduced.

There exists a potential for fruitful research on this subject. One option would be to estimate the costs of hours and days of work using a hedonic wage model. In doing so, we could arrive at estimates of the magnitude to which workers would have to be compensated to vary their hours/days combinations. This would give a good indication of the relative daily and hourly costs of employment. A second option would be to use panel data to analyze the shocks to individual utility functions that result in changes in hours and days for both job-stayers and job-changers. Disaggregating weekly labor supply into its days and hours components is an important, albeit first, step in analyzing a rich variety of policy questions.

Notes

1. Hamermesh (1996) notes that most surveys ask the question, "How many hours did you work last week?"
2. Cogan (1980, 1981), Hanoch (1980b), and Hausman (1980) have all shown that higher fixed costs of entering the labor force result in lower participation rates. Since average fixed costs decline over the number of hours worked, a person must be able to work a minimum number of hours to recoup these costs.
3. Specifically, the questions asked were 1) How many weeks per month did [the subject] usually work at this job? 2) In those weeks, how many paid days per week did he/she usually work? 3) On those days, how many paid hours per day did he/she usually work?
4. In Table 1, the sample is broadened to include the self-employed, but only where appropriate data is available. Unfortunately, the LMAS only includes data on hours and days for a subsample of the self-employed. For this reason, the analysis past Table 1 will be limited to paid workers only.
5. Specifically, we began with a with a sample of 30,924 males and 32,092 females. By eliminating those who held no jobs in 1990, the male (female) sample was reduced by 4,766 (10,326). The self-employed were also dropped (5,131 males and 2,728 females), as were full-time students (2,985 males and 2,746 females). Also eliminated were 179 males and 124 females who did not meet our age criteria, and 70 males and 71 females who claimed to work more than 18 hours per day or earn less than $1.00 per hour at their first job. Another 737 males and 536 females were dropped because they held more than two jobs in 1990. An additional 821 males and 924 females did not meet our criterion of no job overlap. Finally, 5 males (2 females) either earned less than $1.00 per hour or worked more than 18 hours per day at their second jobs (if they held second jobs) and were dropped. This leaves us with 16,280 males and 14,635 females.
6. An additional 140 males and 72 females who held paid jobs preceding self-employment were removed.
7. Direct comparability is a problem because Hamermesh uses an 8-hour day as a standard workday, whereas we define the range 7.5 to 8.5 hours to be standard.
8. The hourly wage rate is not included as a regressor because it is generally derived from earnings per time period and the number of hours worked per time period (the exception is for workers paid by the hour because hourly wage data was collected independently of hours and days data). Such introduction of the wage into the regressions would result in a negative spurious correlation with the hours and days variables.
9. Weiss (1984) found younger workers more likely to quit than older workers. Farber (1980) and Freeman (1980) discussed the lower quit probabilities of unionists in the United States. Blau and Kahn (1981), Meitzen (1986), and Sicherman (1996) all found a negative correlation between tenure and quit behavior.
10. Part of the reason for these inconsistent results may be as a result of our treatment of days as a continuous variable when in fact it is an integer.

11. Throughout this section we use the statistic

    $$z = (\overline{X}_1 = \overline{X}_2) / \sqrt{(s_1^2 / n_1) + (s_2^2 / n_2)},$$

    where $\overline{X}_1, \overline{X}_2$ are the sample means of the two distributions, s_1^2, s_2^2 are the corresponding estimated variances, and n_1, n_2 are the sample sizes. With large sample sizes, this approximates a normal distribution. For testing differences in estimated variances, we use the statistic

    $$F_{n_1 - 1, n_2 - 1} = s_1^2 / s_2^2,$$

    where the variables have already been defined and $s_1^2 \geq s_2^2$.
12. Those respondents who changed jobs because of a labor dispute, a layoff, a company moving or going out of business (i.e., a plant closure), or a dismissal are considered involuntary movers. Voluntary movers changed jobs because of an illness or disability, personal or family responsibilities, to move to a new residence or return to school, a retirement, a new job, or because of a variety of poor working conditions.
13. All such pairwise comparisons are statistically different (at least the 10 percent) in both means and variances of changes. The exceptions are the means of ln (daily) hours at the first job between male stayers and changers, mean changes in ln (weekly) and ln (daily) hours between male voluntary and involuntary changers, and mean changes in ln (daily) hours between female voluntary and involuntary changers.
14. This analysis suffers from censored data because we are only able to observe hours/days changes for job-changers. Altonji and Paxson (1986) have panel data and can use a "difference-in-difference" approach.
15. These comparisons are not strictly equivalent. Hamermesh uses a "difference-in-difference" approach, comparing changers and stayers. Data limitations prevent us from performing similar calculations.
16. The exception is in the case of females who desired extra hours. Given that the distribution is skewed in favor of both more days and hours, the fact that we can't reject this hypothesis comes as little surprise.

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Table A1 Sample Mean Personal and Job Characteristics of Male and Female Paid Workers, Job-Changers, and Job-Stayers[a]

	Males			Females		
Variable	Full sample	Job-stayers	Job-changers	Full sample	Job-stayers	Job-changers
Usual work schedule						
(hours/day) job 1	8.448	8.435	8.556	7.398	7.387	7.516
(hours/day) job 2	N/A	N/A	8.597	N/A	N/A	7.540
(days/week) job 1	5.012	5.011	5.019	4.625	4.617	4.712
(days/week) job 2	N/A	N/A	5.026	N/A	N/A	4.675
Voluntarily left job 1	N/A	N/A	0.640	N/A	N/A	0.757
Extra monthly hours wanted at job 1	0.098	0.092	0.147	0.115	0.108	0.181
Personal characteristics						
Age (%)						
17–19	0.017	0.014	0.054	0.012	0.010	0.032
20–24	0.090	0.081	0.177	0.088	0.077	0.193
25–34	0.303	0.294	0.386	0.324	0.315	0.407
35–44	0.295	0.300	0.240	0.306	0.313	0.239
45–54	0.187	0.196	0.106	0.188	0.197	0.101
55–64	0.107	0.114	0.038	0.082	0.087	0.027
Children						
Number of kids ages 0–2	0.144	0.142	0.150	0.130	0.132	0.110
Number of kids ages 3–5	0.146	0.145	0.152	0.133	0.133	0.132
Number of kids ages > 5	0.907	0.914	0.845	0.897	0.902	0.838
Marital status (%)						
Married	0.748	0.759	0.633	0.745	0.754	0.659
Single	0.204	0.192	0.322	0.157	0.147	0.250
Other	0.048	0.049	0.044	0.098	0.099	0.091
Relationship to family head (%)						
Head	0.815	0.825	0.711	0.253	0.249	0.290
Spouse	0.058	0.056	0.075	0.664	0.674	0.562
Other	0.127	0.118	0.214	0.083	0.077	0.148

(continued)

Table A1 (continued)

Variable	Males Full sample	Males Job-stayers	Males Job-changers	Females Full sample	Females Job-stayers	Females Job-changers
Education (%)						
Elementary	0.094	0.095	0.088	0.054	0.054	0.047
Some high school	0.224	0.221	0.250	0.180	0.179	0.181
Graduated high school	0.227	0.227	0.230	0.272	0.271	0.285
Some postsecondary	0.093	0.091	0.107	0.099	0.097	0.124
Postsecondary diploma	0.126	0.128	0.109	0.191	0.192	0.175
University degree	0.139	0.142	0.105	0.131	0.134	0.096
Trade	0.097	0.096	0.111	0.075	0.073	0.092
Region (%)						
BC	0.105	0.104	0.115	0.099	0.097	0.113
Prairies	0.269	0.266	0.291	0.289	0.287	0.305
Ontario	0.205	0.208	0.183	0.210	0.210	0.216
Quebec	0.160	0.161	0.152	0.149	0.154	0.105
Atlantic	0.261	0.261	0.259	0.252	0.252	0.260
Native language (%)						
English	0.722	0.720	0.744	0.738	0.733	0.789
French	0.187	0.187	0.184	0.175	0.179	0.144
Other	0.092	0.093	0.072	0.087	0.089	0.067
Other characteristics (%)						
Immigrant	0.097	0.099	0.079	0.093	0.094	0.078
Visible minority	0.032	0.032	0.030	0.034	0.034	0.027
Job characteristics						
Firm size (%)						
19 or fewer employees	0.218	0.203	0.345	0.260	0.254	0.309
20–99 employees	0.156	0.151	0.198	0.149	0.149	0.149
100–499 employees	0.122	0.124	0.106	0.126	0.127	0.115
500 or more employees	0.363	0.379	0.220	0.326	0.333	0.263
Don't know	0.141	0.143	0.131	0.139	0.137	0.165

Variable	Males			Females		
	Full sample	Job-stayers	Job-changers	Full sample	Job-stayers	Job-changers
Industry group[b]						
Goods sector						
Primary (01–08)	0.091	0.089	0.107	0.031	0.031	0.029
Construction (29–30, 52)	0.098	0.089	0.175	0.016	0.015	0.019
Manufacturing (09–28)	0.239	0.244	0.194	0.103	0.102	0.108
Service Sector						
Distributive services (31–35)	0.182	0.187	0.146	0.069	0.069	0.062
Business services (37–39, 44)	0.058	0.057	0.062	0.115	0.112	0.144
Consumer services (36, 43, 45–47)	0.144	0.135	0.214	0.286	0.275	0.386
Education, health and welfare (40–42)	0.095	0.101	0.050	0.303	0.316	0.178
Public administration (48–51)	0.093	0.097	0.052	0.078	0.078	0.074
Occupation group[b]						
Managerial and administrative (01–03)	0.133	0.137	0.087	0.104	0.103	0.112
Professional and technical (04–16)	0.128	0.131	0.106	0.225	0.234	0.143
Clerical (17–22)	0.056	0.056	0.061	0.308	0.309	0.309
Sales (23–24)	0.059	0.057	0.077	0.086	0.085	0.098
Service (25–28)	0.084	0.084	0.087	0.167	0.163	0.203
Primary (29–33)	0.064	0.062	0.087	0.018	0.018	0.022
Processing (34-35)	0.063	0.064	0.050	0.034	0.032	0.050
Machining, fabricating, assembling, and repairing (36–42)	0.161	0.163	0.141	0.030	0.031	0.029

(continued)

Table A1 (continued)

Variable	Males Full sample	Males Job-stayers	Males Job-changers	Females Full sample	Females Job-stayers	Females Job-changers
Construction trades (43–45)	0.124	0.117	0.187	0.004	0.003	0.011
Transport operating and materials handling (46–49)	0.126	0.127	0.116	0.022	0.022	0.024
Other occupations (50)	0.000	0.000	0.001	0.000	0.000	0.000
Other						
Covered by union agreement (%)	0.453	0.476	0.266	0.384	0.402	0.214
Part-time employment (<120 hours per month) (%)	0.043	0.038	0.086	0.246	0.248	0.222
Tenure at job (weeks)	412	445	136	296	315	114
Covered by pension (%)	0.396	0.395	0.409	0.328	0.327	0.336
Number of observations	16,280	14,577	1,563	14,635	13,245	1,318

[a] Values may not add to 100% due to rounding.

[b] LMAS industry codes.

Table A2 Full Regression Results from Table 2[a,b]

Independent variable	Males						Females					
	No job controls			Job controls			No job controls			Job controls		
	Hours	Days	*t*	Hours	Days	*t*	Hours	Days	*t*	Hours	Days	*t*
Personal characteristics												
Age												
20–24 years	0.0493	0.0385		0.0501	0.0413		0.0327	0.0853	***	0.0248	0.0787	***
	(4.067)	(3.667)		(4.239)	(3.970)		(1.441)	(3.373)		(1.115)	(3.160)	
25–34 years	0.0549	0.0412		0.0537	0.0460		0.0225	0.0785	***	0.0010	0.0553	***
	(4.516)	(3.918)		(4.522)	(4.392)		(1.008)	(3.156)		(0.045)	(2.251)	
35–44 years	0.0371	0.0322		0.0427	0.0392		0.0106	0.0688	***	–0.0175	0.0385	***
	(3.006)	(3.007)		(3.518)	(3.667)		(0.467)	(2.735)		(0.787)	(1.542)	
45–54 years	0.0227	0.0310		0.0317	0.0377		–0.0176	0.0395	***	–0.0413	0.0118	***
	(1.792)	(2.832)		(2.530)	(3.414)		(0.763)	(1.541)		(1.816)	(0.463)	
55–64 years	–0.0165	0.0198	**	–0.0036	0.0269	*	–0.0807	–0.0362		–0.1025	–0.0637	
	(1.251)	(1.732)		(0.278)	(2.337)		(3.371)	(1.360)		(4.319)	(2.399)	
Children												
Number of kids ages 0–2	–0.0023	–0.0031		–0.0030	–0.0035		–0.0195	–0.0742	*	–0.0163	–0.0681	*
	(0.579)	(0.893)		(0.752)	(1.003)		(2.867)	(9.788)		(2.430)	(9.102)	
Number of kids ages 3–5	0.0059	0.0005		0.0060	0.0012		–0.0332	–0.0645	*	–0.0320	–0.0616	*
	(1.507)	(0.152)		(1.590)	(0.354)		(5.153)	(9.012)		(5.078)	(8.726)	

Number of kids ages > 5	0.0022	0.0007		0.0011	0.0005		–0.0192	–0.0277	*	–0.0167	–0.0224	***
	(1.445)	(0.559)		(0.725)	(0.348)		(7.749)	(10.082)		(6.843)	(8.206)	
Marital status												
Single	–0.0216	–0.0094	***	–0.0170	–0.0070		0.0062	0.0306	***	0.0069	0.0317	***
	(3.966)	(2.002)		(3.203)	(1.489)		(0.586)	(2.583)		(0.660)	(2.717)	
Other	0.0033	–0.0106		0.0031	–0.0104		0.0182	0.0395		0.0192	0.0400	
	(0.477)	(1.756)		(0.459)	(1.737)		(1.648)	(3.215)		(1.779)	(3.304)	
Relationship to family head												
Spouse	–0.0070	–0.0106		–0.0018	–0.0088		–0.0201	–0.0207		–0.0195	–0.0191	
	(1.115)	(1.951)		(0.294)	(1.637)		(2.307)	(2.136)		(2.281)	(1.995)	
Other	–0.0293	–0.0147	***	–0.0309	–0.0168	***	–0.0063	–0.0229		–0.0042	–0.0217	
	(4.697)	(2.715)		(5.087)	(3.144)		(0.602)	(1.956)		(0.408)	(1.873)	
Education												
Some high school	–0.0149	–0.0168		–0.0047	–0.0140		0.0117	–0.0229	**	0.0183	–0.0183	**
	(2.626)	(3.401)		(0.847)	(2.849)		(1.033)	(1.808)		(1.621)	(1.448)	
Graduated high school	–0.0271	–0.0253		–0.0048	–0.0205	**	0.0215	–0.0239	*	0.0233	–0.0292	*
	(4.671)	(5.028)		(0.826)	(4.027)		(1.935)	(1.933)		(2.046)	(2.287)	
Some postsecondary	–0.0410	–0.0254	***	–0.0128	–0.0208		0.0225	–0.0341	*	0.0195	–0.0393	*
	(5.996)	(4.284)		(1.867)	(3.440)		(1.787)	(2.440)		(1.509)	(2.720)	
Postsecondary diploma	–0.0524	–0.0194	*	–0.0204	–0.0167		0.0603	–0.0586	*	0.0479	–0.0475	*
	(8.169)	(3.499)		(3.119)	(2.900)		(5.272)	(4.605)		(3.954)	(3.507)	
University degree	–0.0629	–0.0173	*	–0.0098	–0.0171		0.0513	–0.0294	*	0.0227	–0.0076	***
	(10.062)	(3.189)		(1.383)	(2.738)		(4.274)	(2.200)		(1.733)	(0.518)	

(continued)

Table A2 (continued)

Independent variable	Males						Females					
	No job controls			Job controls			No job controls			Job controls		
	Hours	Days	*t*	Hours	Days	*t*	Hours	Days	*t*	Hours	Days	*t*
Trade	–0.0243	–0.0178		–0.0037	–0.0147		0.0343	–0.0250	*	0.0407	–0.0193	*
	(3.628)	(3.058)		(0.551)	(2.518)		(2.609)	(1.713)		(3.040)	(1.290)	
Region												
BC	0.0140	–0.0104	*	0.0099	–0.0129	*	–0.0104	–0.0324	***	–0.0010	–0.0201	
	(2.568)	(2.202)		(1.851)	(2.729)		(1.176)	(3.296)		(0.109)	(2.056)	
Prairies	0.0113	0.0035		0.0080	–0.0029	**	–0.0154	–0.0365	**	–0.0041	–0.0274	**
	(2.649)	(0.948)		(1.886)	(0.780)		(2.331)	(4.965)		(0.626)	(3.732)	
Quebec	–0.0279	–0.0022	*	–0.0183	–0.0012	**	–0.0300	–0.0021	**	–0.0294	–0.0029	***
	(4.457)	(0.408)		(2.986)	(0.227)		(3.006)	(0.192)		(3.005)	(0.267)	
Atlantic	0.0152	0.0133		0.0134	0.0080		0.0172	0.0255		0.0257	0.0335	
	(3.440)	(3.469)		(3.043)	(2.071)		(2.459)	(3.284)		(3.672)	(4.282)	
Native language												
French	0.0036	–0.0041		–0.0021	–0.0034		–0.0024	0.0022		–0.0025	0.0065	
	(0.678)	(0.888)		(0.399)	(0.760)		(0.287)	(0.232)		(0.297)	(0.700)	
Other	–0.0152	–0.0042		–0.0191	–0.0030	**	–0.0063	0.0015		–0.0084	–0.0005	
	(2.377)	(0.758)		(3.068)	(0.551)		(0.640)	(0.132)		(0.870)	(0.049)	
Other												
Immigrant	0.0095	0.0104		0.0115	0.0100		–0.0019	0.0056		–0.0015	0.0092	
	(1.498)	(1.906)		(1.870)	(1.834)		(0.193)	(0.515)		(0.155)	(0.851)	

Visible minority	–0.0243	–0.0059	***	–0.0153	–0.0025		0.0475	0.0765		0.0427	0.0680	
	(2.767)	(0.781)		(1.794)	(0.328)		(3.446)	(4.994)		(3.159)	(4.498)	
Job characteristics												
Firm size												
20–99 employees				0.0145	–0.0107	*				0.0509	0.0298	**
				(3.041)	(2.547)					(6.797)	(3.554)	
100–499 employees				0.0209	–0.0137	*				0.0515	0.0326	
				(3.876)	(2.899)					(6.197)	(3.509)	
500 or more employees				0.0147	–0.0188	*				0.0287	0.0329	
				(3.278)	(4.756)					(4.298)	(4.406)	
Don't know				0.0135	–0.0022	**				–0.0048	0.0236	*
				(2.671)	(0.488)					(0.629)	(2.746)	
Industry group												
Goods sector												
Construction				–0.0460	0.0143	*				–0.0730	0.0229	*
				(5.366)	(1.892)					(2.893)	(0.812)	
Manufacturing				–0.0808	–0.0040	*				0.0266	0.0477	
				(11.254)	(0.629)					(1.373)	(2.198)	
Service sector												
Distributive services				–0.0926	0.0110	*				–0.0285	0.0270	**
				(12.776)	(1.719)					(1.486)	(1.257)	
Business services				–0.1093	–0.0003	*				–0.0171	0.0285	***
				(12.354)	(0.041)					(0.925)	(1.378)	

(continued)

Table A2 (continued)

Independent variable	Males						Females					
	No job controls			Job controls			No job controls			Job controls		
	Hours	Days	*t*	Hours	Days	*t*	Hours	Days	*t*	Hours	Days	*t*
Consumer services				–0.1133	0.0016	*				–0.0197	–0.0059	
				(14.890)	(0.232)					(1.106)	(0.296)	
Education, health and welfare				–0.1377	–0.0007	*				–0.0723	–0.0310	
				(16.364)	(0.100)					(3.979)	(1.524)	
Public administration				–0.1124	–0.0012	*				–0.0304	0.0243	**
				(14.102)	(0.165)					(1.588)	(1.134)	
Occupation group												
Managerial and administrative				0.0378	0.0389					0.0768	0.0610	
				(5.207)	(6.086)					(9.369)	(6.649)	
Professional and technical				0.0219	0.0312					0.0717	–0.0427	*
				(2.821)	(4.575)					(8.896)	(4.731)	
Sales				0.0321	0.0268					–0.0033	–0.0409	*
				(3.835)	(3.631)					(0.359)	(3.960)	
Service				0.0283	–0.0026	*				–0.0265	–0.0163	
				(3.599)	(0.375)					(3.442)	(1.899)	
Primary				0.0563	0.0678					0.0521	0.0966	
				(5.757)	(7.874)					(2.316)	(3.837)	
Processing				0.0663	0.0079	*				0.0668	0.0174	**
				(7.656)	(1.041)					(4.277)	(0.995)	

Machining, fabricating, assembling and repairing				0.0320	0.0269					0.0706	0.0172	**
				(4.522)	(4.305)					(4.524)	(0.983)	
Construction trades				0.0430	0.0236	***				0.1711	0.1125	
				(5.326)	(3.316)					(4.468)	(2.625)	
Transport operating and materials handling				0.0698	0.0153					–0.0185	–0.0136	
				(9.613)	(2.388)					(1.150)	(0.754)	
Other occupation				0.1344	0.0720					–0.0233	–0.0521	
				(1.981)	(1.206)					(0.211)	(0.420)	
Other												
Covered by union agreement				–0.0161	–0.0184					0.0313	0.0181	
				(4.735)	(6.152)					(5.281)	(2.730)	
Tenure at job (weeks/100)				–0.0001	0.0007	***				0.0026	0.0053	**
				(0.304)	(2.152)					(3.534)	(6.316)	
Covered by pension				–0.0024	0.0002					0.0105	0.0096	
				(0.849)	(0.080)					(2.190)	(1.789)	
Constant	2.1163	1.5883	*	2.1341	1.5712	*	1.9751	1.5201	*	1.9535	1.4841	*
	(151.034)	(130.867)		(127.509)	(106.577)		(75.150)	(52.004)		(61.491)	(41.755)	
R^2	0.0313	0.0086		0.0841	0.0305		0.0276	0.0696		0.0408	0.0713	

(continued)

Table A2 (continued)

Independent variable	Males						Females					
	No job controls			Job controls			No job controls			Job controls		
	Hours	Days	*t*	Hours	Days	*t*	Hours	Days	*t*	Hours	Days	*t*
Correlation coefficient of residuals	0.1455			0.1379			0.1545			0.1371		
Breusch–Pagan test of independence (*p*–value)	0.0000			0.0000			0.0000			0.0000		
Number of observations	16,280			16,280			14,635			14,635		

[a] *, **, and *** denote that the pairwise *t*-statistics are significant at 1%, 5%, and 10%, respectively.

[b] Omitted categorical variables are age 17–19, married, head of household, elementary education, Ontario, English, 19 or fewer employees, primary industry group and clerical occupation group.

Table A3 Full Probit Results from Table 3[a,b]

	Males				Females			
	No controls		With controls		No controls		With controls	
Independent variable	Coefficient	Partial deriv.	Coefficient	Partial deriv.	Coefficient	Partial deriv.	Coefficient	Partial deriv.
ln (hours per day)	0.2869 (3.442)	0.048	0.0806 (0.890)	0.010	0.3203 (4.394)	0.051	0.2358 (2.997)	0.027
ln (days per week)	0.1946 (1.832)	0.033	–0.0321 (0.296)	–0.004	0.3093 (4.522)	0.049	0.2330 (3.188)	0.027
Extra hours wanted	1.7816 (4.231)	0.548	0.5529 (1.263)	0.092	0.8750 (2.913)	0.208	0.5999 (1.919)	0.098
Extra hours wanted*ln (hours)	–0.2810 (1.515)	–0.047	0.0388 (0.199)	0.005	–0.1086 (0.791)	–0.017	–0.0309 (0.213)	–0.004
Extra hours wanted*ln (days)	–0.5910 (3.518)	–0.100	–0.3389 (1.950)	–0.041	–0.1856 (1.651)	–0.029	–0.2079 (1.763)	–0.024
Personal characteristics								
Age								
20–24 years			–0.3188 (3.487)	–0.031			–0.0973 (0.836)	–0.011
25–34 years			–0.4625 (4.925)	–0.049			–0.2862 (2.459)	–0.031

(continued)

Table A3 (continued)

	Males				Females			
	No controls		With controls		No controls		With controls	
Independent variable	Coefficient	Partial deriv.	Coefficient	Partial deriv.	Coefficient	Partial deriv.	Coefficient	Partial deriv.
35–44 years			–0.5028 (5.145)	–0.052			–0.4378 (3.641)	–0.045
45–54 years			–0.6083 (5.852)	–0.055			–0.5785 (4.616)	–0.051
55–64 years			–0.8859 (7.612)	–0.063			–0.7840 (5.477)	–0.055
Children								
Number of kids ages 0–2			–0.1106 (2.771)	–0.013			–0.2743 (5.783)	–0.032
Number of kids ages 3–5			–0.0032 (0.085)	0.000			–0.0619 (1.425)	–0.007
Number of kids ages > 5			–0.0196 (1.242)	–0.002			–0.0345 (2.015)	–0.004
Marital status								
Single			–0.0396 (0.769)	–0.005			–0.0186 (0.275)	–0.002
Other			–0.0157 (0.212)	–0.002			0.0413 (0.553)	0.005

Relationship to family head				
Spouse	0.1353	0.018	–0.0654	–0.008
	(2.262)		(1.136)	
Other	–0.0426	–0.005	–0.0045	–0.001
	(0.750)		(0.070)	
Education				
Some high school	0.0086	0.001	0.0349	0.004
	(0.147)		(0.427)	
Graduated high school	–0.0221	–0.003	0.0516	0.006
	(0.360)		(0.627)	
Some postsecondary	0.0860	0.011	0.1434	0.018
	(1.210)		(1.581)	
Postsecondary diploma	0.0010	0.000	0.1432	0.018
	(0.014)		(1.637)	
University degree	0.0347	0.004	0.1443	0.018
	(0.453)		(1.501)	
Trade	0.0977	0.012	0.1749	0.023
	(1.415)		(1.859)	
Region				
BC	0.0528	0.007	0.0409	0.005
	(0.942)		(0.691)	
Prairies	0.0300	0.004	0.0064	0.001
	(0.672)		(0.142)	

(continued)

Table A3 (continued)

	Males				Females			
	No controls		With controls		No controls		With controls	
Independent variable	Coefficient	Partial deriv.	Coefficient	Partial deriv.	Coefficient	Partial deriv.	Coefficient	Partial deriv.
Quebec			–0.0626	–0.007			–0.2722	–0.027
			(0.962)				(3.877)	
Atlantic			–0.0807	–0.009			–0.1201	–0.013
			(1.703)				(2.453)	
Mother tongue								
French			0.0293	0.004			0.0802	0.010
			(0.548)				(1.397)	
Other			–0.0422	–0.005			–0.0693	–0.008
			(0.614)				(0.962)	
Other								
Immigrant			0.0136	0.002			0.0020	0.000
			(0.202)				(0.028)	
Visible minority			–0.1126	–0.012			–0.1648	–0.017
			(1.225)				(1.683)	
Job characteristics								
Firm size								
20–99 employees			0.0073	0.001			–0.0426	–0.005
			(0.163)				(0.839)	

100–499 employees	–0.0458	–0.005	0.0668	0.008
	(0.844)		(1.160)	
500 or more employees	–0.1421	–0.017	–0.0138	–0.002
	(3.131)		(0.300)	
Don't know	–0.0832	–0.010	0.0510	0.006
	(1.667)		(1.005)	
Industry group				
Goods sector				
Construction	0.0632	0.008	0.1419	0.018
	(0.738)		(0.784)	
Manufacturing	–0.0402	–0.005	0.1601	0.020
	(0.528)		(1.151)	
Service sector				
Distributive services	–0.0682	–0.008	0.2621	0.036
	(0.881)		(1.892)	
Business services	–0.0597	–0.007	0.3302	0.047
	(0.656)		(2.493)	
Consumer services	0.0228	0.003	0.2928	0.038
	(0.292)		(2.273)	
Education, health and welfare	–0.2051	–0.022	0.1318	0.016
	(2.178)		(0.998)	
Public administration	–0.1125	–0.013	0.3878	0.058
	(1.268)		(2.813)	

(continued)

Table A3 (continued)

	Males				Females			
	No controls		With controls		No controls		With controls	
Independent variable	Coefficient	Partial deriv.	Coefficient	Partial deriv.	Coefficient	Partial deriv.	Coefficient	Partial deriv.
Occupation group								
Managerial and administrative			–0.2265 (2.878)	–0.024			0.0379 (0.670)	0.005
Professional and technical			–0.0082 (0.100)	–0.001			–0.0089 (0.151)	–0.001
Sales			–0.0788 (0.942)	–0.009			–0.0519 (0.846)	–0.006
Service			–0.0837 (1.028)	–0.010			0.0183 (0.353)	0.002
Primary			–0.1227 (1.222)	–0.014			0.1909 (1.248)	0.025
Processing			–0.1824 (1.942)	–0.019			0.2698 (2.685)	0.038
Machining, fabricating, assembling and repairing			–0.1924 (2.573)	–0.021			0.1031 (0.959)	0.013
Construction trades			–0.0147 (0.174)	–0.002			0.6209 (2.968)	0.112
Transport operating and materials handling			–0.1643 (2.130)	–0.018			0.0189 (0.174)	0.002

Other occupation			−0.0703 (0.110)	−0.008			variable dropped	
Other								
Covered by union agreement			−0.2124 (5.768)	−0.025			−0.2854 (6.611)	−0.032
Tenure at job (weeks/100)			−0.1179 (18.130)	−0.014			−0.1460 (15.541)	−0.017
Covered by pension			0.0185 (0.621)	0.002			0.0153 (0.459)	0.002
Constant	−2.2614 (9.773)		−0.3095 (1.096)		−2.5003 (14.577)		−1.6369 (6.028)	
Pseudo R^2	0.007		0.124		0.012		0.119	
χ^2 (5, 55, 5 and 54 d.f.)	69.32		1277.03		110.18		1055.43	
Observed *P*		0.096		0.096		0.090		0.090
Predicted *P*		0.095		0.061		0.087		0.058
Number of observations	16,280		16,280		14,635		14,629	

[a] Absolute *t*-ratios are in parentheses.

[b] Omitted categorical variables are age 17–19, married, head of household, elementary education, Ontario, English native language, 19 or fewer employees, primary industry and clerical occupation.

8

Hours Constraints

Theory, Evidence, and Policy Implications

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In response to persistent unemployment, particularly in Europe, there have been calls to reduce the length of the workweek in order to share the available work more equally. Implicitly, advocates of these mandated hours reductions believe that the demand for hours of work is inelastic and independent of the number of workers used to fill those hours. Therefore overall employment can be increased by reducing the number of hours that each individual works, a policy often referred to as work sharing. The view that modern workers would actually like to reduce their work hours adds to the attractiveness of this proposal. Juliet Schor's book *The Overworked American* proved enormously popular. Despite stinging criticism from academic economists, the book appeals to professional women, many of whom are in dual-career families and feel caught between the high demands of their jobs and their families.

The Canadian "Report of the Advisory Group on Working Time and the Distribution of Work" (1994) takes a more cautious approach. It suggests that about half of sustained reductions in overtime eventually are translated into new jobs. On the basis of this and other arguments, it recommends reducing the legislated standard workweek to 40 hours in those provinces where the legislated standard exceeds 40. In addition, it recommends giving employees the right to refuse to work more than 40 hours per week. Finally, it recommends allowing a maximum of 100 hours of compensated overtime per year. Additional overtime would have to be offset by reduced hours at other times. In effect, it therefore recommends a legislated maximum average workweek of 42 hours. We note in passing that this would be the maximum time spent per week on a *single* job. Many people hold more than one job. Such legislation would undoubtedly increase these numbers.

It is relatively easy for economists to dismiss both the calls for mandated hours reductions and Schor's book. In a simple model of hours determination, hours are set optimally. Any interference with the market must reduce welfare. Indeed, it is trivial to produce examples in which reducing the workweek actually reduces employment due to decreased efficiency, suggesting that there even might not be an efficiency/equity trade-off.

Nevertheless, there are deficiencies in this simple economic argument. The models which are used implicitly or explicitly to show that reducing the workweek need not increase employment and therefore may not reduce unemployment are models in which the labor market clears, employment is efficient and there is no unemployment. Having assumed away unemployment, it is difficult to see how we can evaluate programs designed to reduce it. Clearly, more economic analysis is called for.

The recommendation that the workweek should be reduced contrasts with many workers' perceptions that they are underemployed. Most workers do not want to reduce work hours in return for a proportionate pay reduction. The overwhelming evidence from both the United States and Canada is that far more workers would like to *increase* their hours than would like to decrease them, although the European evidence is more mixed. Thus, where some see mandated hours reductions lowering unemployment, others see it exacerbating underemployment.

In this chapter, we do not claim to resolve the issue of whether mandated hours reductions are a viable mechanism for reducing unemployment; our goals are more modest. We consider, both empirically and theoretically, workers' survey responses regarding how many hours they would prefer to work at the same *hourly* rate compared to how much they actually do work. When preferred hours diverge from actual hours, hours are *constrained*.

We have two objectives in examining hours constraints. The first is to assess whether hours constraints are indicative of some sort of problem in the labor market, particularly one of underemployment. The second is to use the information on hours constraints to further our understanding of the functioning of the labor market and determination of working hours.

We conclude that in the United States and Canada, the direction of hours constraints is clear: wanting to work additional hours is more prevalent than wanting to work fewer hours. The evidence is less clear for Europe. The most promising avenue for explaining hours constraints is the development of models of imperfect matching in the labor market, possibly supplemented by issues associated with long-term contracting. We develop a simple example, a model of bilateral search, where although most workers want to work fewer hours, imposing a legally mandated shorter workweek could worsen unemployment and reduce the well-being of workers. Thus we urge caution regarding proposals to promote work sharing by requiring a shorter workweek—the data suggest that more people would prefer to work more hours than fewer hours, and even if people preferred fewer hours, we cannot identify welfare-improving regulation without understanding why both hours constraints and unemployment exist.

THE EXTENT OF HOURS CONSTRAINTS

Survey research on whether people in the United States would like to work more, fewer, or the same number of hours dates back to at least 1966. Since then, five surveys have asked comparable questions regarding desired work hours.[1] Table 1 gives the results of these surveys. All five surveys reveal the same general tendency: more than 40 percent of respondents would like to change their hours of work. Of these, a clear majority want more work not less. Differences in the samples make it difficult to draw strong conclusions about trends.

The Panel Study of Income Dynamics (PSID) also monitored the relation between desired and actual hours through 1987. In the PSID, workers were first asked whether more hours were available on any of their jobs. Those who could not have worked more were then asked if they would have liked to work more. Similarly, they were asked if they could have worked less and, if not, whether they would have liked to work less. One weakness of the PSID is that some salaried workers responded that they could have worked more but, in a subsequent question, revealed that they would not have been paid for the work. Nevertheless, these workers were not asked if they would have liked to have worked more in return for more pay.

Table 1 Selected Survey Results on U.S. Workers' Desire to Work More, Fewer, or the Same Hours (%)

Year	More	Fewer	Same
1966	34	10	56
1978	28	11	61
1985	28	8	65
1991	33	6	62
1995	26	14	55

SOURCE: 1966: George Katona and his associates, sample of household heads (Katona et al. 1971). The exact question was, "Some people would like to work more hours a week, if they could be paid for it. Others would prefer to work fewer hours per week even if they earned less. How do you feel about this?"
1978: Conducted by Louis Harris Associates, sample of employed civilians, aged 17 and over. The question was, "If you had a choice, would you prefer to work the same number of hours and earn the same money, fewer hours at the same rate of pay and earn less money, or more hours at the same rate of pay and earn more money?"
1985: Current Population Survey supplement, sample of employed persons, aged 17 and over (Shank 1986). Question identical to the one above.
1991: International Social Survey Programme. This survey asked, "Think of the number of hours you work and the money you earn in your main job, including regular overtime. If you only had one of these three choices, which of the following would you prefer: work longer hours and earn more money; work the same number of hours and earn the same money; work fewer hours and earn less money?" (Bell and Freeman 1995).
1995: The Gallup Poll, sample of employed persons, aged 18 and over (*USA Today*, April 10, 1995). The exact question was, "If you could, which of the following situations would you choose: fewer hours on the job but less income, the same number of hours and income that you now have, or more hours on the job and more income?"

Despite this weakness, the PSID has the advantage of offering a consistent time series. Moreover, work by Ham (1982) and Altonji and Paxson (1988) shows that the responses to the constraints questions have predictive power for behavior. In addition, Kahn and Lang (1992) show that for wage earners, the PSID questions give results that are similar to those obtained using the questions in the other surveys summarized in Table 1.

Table 2 presents the fractions of PSID respondents who say they would have liked to work more or fewer hours over the period 1968–1987. There may be some bias in time trends in these PSID figures.

The PSID, limited to household heads, follows families through time. While break-off families are added to the sample, there is a risk that part of observed time trends captures changes in household heads over the life cycle. Additionally, in the early years, the low-income population was oversampled. However, over time, regression toward the mean in earnings has led to progressively less oversampling of the low-income population. If these two factors introduce a bias, it should push us toward finding a reduction in the desire to work more, because our work shows that older workers are less likely and poorer workers more likely to want more work.

Table 2 Desire for Different Work Hours (U.S.) 1968–1987 (%)

	Men		Women	
Year	Fewer	More	Fewer	More
1968	6	14	3	13
1969	6	20	6	17
1970	3	20	5	20
1971	6	20	6	21
1972	6	20	6	19
1973	6	20	6	18
1974	5	20	6	24
1975	5	18	5	23
1976	5	20	4	24
1977	9	22	7	18
1978	4	24	7	18
1979	6	22	7	18
1980	5	21	6	20
1981	7	23	5	23
1982	6	26	6	25
1983	7	26	5	23
1984	8	22	6	23
1985	8	22	7	21
1986	6	26	6	21
1987	6	21	6	22

SOURCE: Panel Study of Income Dynamics.

In fact, there is no clear trend in the data. The 1968 numbers are quite different from those of later years. Ignoring this first year of the survey generates a positive correlation between time and both wanting to work more and wanting to work less for women. For men, there is a positive time trend for wanting to work more, but it is significant at only the 0.1 level.

Despite the similarity of these responses to questions intended to measure the same phenomenon, there are other ways of framing the question that generate different answers. A 1978 survey (Best 1981) asked workers what the largest portion of their current yearly income they would be willing to give up for shorter workdays (shorter workweeks, more vacation). The options offered were designed to involve proportional cuts in pay (e.g., "2 percent [1/50] of your income for 10 minutes off each workday), although given variation in the length of the workday and number of days worked each year, these may not have been exactly right. Nevertheless, 23 percent said they would take a pay cut for a shorter day, 26 percent for a shorter week, and 42 percent for more vacation. Note, however, that no similar question inquired about possible increases in hours for additional income.

The Canadian Survey of Work Reduction (SWR), conducted in 1986, is particularly helpful for looking at the impact of question wording. While only 17.3 percent of Canadians responded that they would take a pay cut in return for more time off, 26.7 percent were willing to forego some or all of their anticipated pay increase for more time off.[2]

The Advisory Group on Working Time (1994) reports Benimadhu's (1987) calculations from the SWR that 30.7 percent of Canadian workers preferred fewer working hours while 32.1 percent preferred longer working hours (p. 87). The Advisory Group concludes that "[t]he survey strikingly captures Canadians' 'indeterminate' mood regarding working time . . . ." (p. 25).

There are reasons for being skeptical about this conclusion. First, the calculation of people preferring fewer hours included all workers who answered either of the two questions positively (i.e., whether they would be willing to take a pay cut for time off and whether they would be willing to forego part of their pay increase for time off). In contrast, there was only a single question asking whether people preferred to work more hours for more pay (see note 2).

Second, many respondents give contradictory answers to these three different survey questions about hours constraints. In fact, only 12 percent of respondents answer all three questions in a manner consistent with wanting reduced hours.[3] Twenty-seven percent, more than twice as many, answer all three questions in a manner consistent with wanting additional work hours.

Third, respondents who "expressed an interest" in working less were asked why they were interested in less work. Almost half rated as "very important" at least one of various responses that are inconsistent with a true preference for less work: giving others a chance for work, avoiding being laid off, starting a business, looking for other work, running an existing business, or working at a second job. In fact, more than one-quarter rated avoiding being laid off as a "very important" reason for wanting to work less, and more than half rated this as at least "somewhat important." We must therefore exercise extreme caution in interpreting the SWR as revealing a desire for more leisure among a large number of Canadians.

In a similar vein, people who respond that they would like more work might actually want to cut hours at a second job but not work additional total hours. However, so few of these respondents work at a second job that excluding them would not change our estimates of wanting more work.

Given the difficulties of interpreting the responses to the questions regarding the desire for less work, and given the fact that the survey does not inquire why individuals respond that they want to work more, we must be somewhat guarded in our assessment of the results of the survey. Nevertheless, it seems to us that the evidence suggests that Canadians are far more likely to want to work additional hours than to work fewer hours.

This view is reinforced by the results of the 1995 Survey of Work Arrangements (Drolet and Morissette 1997), which asked, "At this job, given the choice, would . . . , at his/her current wage rate, prefer to work: 1) fewer hours for less pay, 2) more hours for more pay, 3) the same hours for the same pay." The survey found that 27 percent of Canadians preferred, at their current wage rate, more hours for more pay, compared with only 6 percent who preferred fewer hours for less pay. These results are quite close to those obtained for similar questions in the United States.

Similar questions have been asked in other countries. Unfortunately, the results of two major surveys, the first conducted by the International Social Survey Program (ISSP) and the second by the European Union, conflict quite sharply. Table 3 gives the results of the ISSP survey, conducted in 1989. Respondents were asked, "If you had a choice, would you prefer to work: 1) the same number of hours and earn the same money, 2) fewer hours at the same rate of pay and earn less money, or 3) more hours at the same rate of pay and earn more money?"

In Table 3, the United States looks similar to other OECD countries. It has a relatively high fraction of workers who want to work more, but this proportion is not substantially higher than in Ireland and Italy. Similarly, relatively few people in the United States want to work less, but that is true of most other OECD countries. In every country more people want to work more than less, although the difference is not large in Germany.

Table 4 gives responses to the European Union survey, also conducted in 1989. The survey asked, "Assuming that your hourly rate remained unchanged, would you like to work less, as long, or longer?" While this question does not appear to be significantly different from the question used by the ISSP, the survey results are dramatically different. A large minority of workers in all countries reply that they would like to work less. In all the European Union countries, the fraction wanting less work is significantly higher than the fraction wanting more.

Table 3 Desire for Different Work Hours in Various Countries

Country	More	Fewer	Same
Austria	23	8	68
Germany	14	10	76
Ireland	30	5	65
Italy	31	7	62
Netherlands	18	12	70
Northern Ireland	27	6	68
Norway	24	7	69
United Kingdom	24	8	68
United States	33	6	62

SOURCE: Bell and Freeman (1995).

Table 4 Desire for Different Work Hours in European Union

Country	More	Less	Same	Part-time, want full-time
Belgium	7	28	43	2
Denmark	9	29	61	1
France	9	39	52	10
Germany	4	38	55	1
Greece	15	28	57	11
Ireland	11	18	65	–
Italy	8	39	50	16
Netherlands	8	31	56	2
Portugal	2	49	46	10
Spain	12	42	44	15
United Kingdom	12	33	50	2

SOURCE: Commission of the European Communities (1991, Tables 22, 23).
The first three columns give responses to the question, "Assuming that your hourly rate remained unchanged, would you like to work less, as long, or longer?"
The last column gives the percentage of all workers who are *both* part-time and answer "yes" in response to the question, "Would you rather have full-time employment?"

We note some other results in the European Union survey that make the results in Table 4 even more surprising. The survey also asked part-time workers whether they would prefer full-time work; results are given in the last column of Table 4. In France, Italy, Portugal, and Spain, more people are part-time and want full-time work than say that they would prefer to work more. While this is technically feasible (e.g., if part-timers respond that they want full-time work because full-time work is compensated at a higher hourly rate than part-time work), the counterintuitive result is concerning.

Another surprising aspect of the European Union study is the difference between answers to the question about wanting more or fewer hours of work at their present hourly rate and answers to a question regarding willingness to trade pay raises for shorter hours. The correlation across countries of the percentage wanting fewer hours in the two questions is only 0.05, although the average across countries is not very dissimilar (34 percent versus 30 percent).

While there is no formal contradiction among the different answers to the different questions, we find these differences disturbing. However, the high number of Europeans desiring shorter hours in the European Union survey seems corroborated by the British Household Panel Survey (BHPS) of 1991, which asked, "Thinking about the hours you work, assuming that you would earn the same amount per hour as at present, would you prefer to: work *fewer* hours than you do now; work *more* hours than you do now; or carry on working the *same number* of hours?" Among male employees age 21–64, 36 percent respond fewer, 7 percent more, and 56 percent the same (Stewart and Swaffield 1995). These results are quite similar to the European Union survey. We find the face validity of the BHPS to be the greatest of the three surveys, because it seems to make it clear that the hourly rate would be unchanged. The contradictions within the European Union survey and between the European Union and ISSP surveys remain a matter of concern.

Finally we note that older European surveys indicated that wanting more work is more common than wanting less work. Katona et al. (1971) report the answers to the question, "Some people would like to work more hours per week if they could be paid for it. Others would prefer to work fewer hours per week even if they earned less. How do you feel about this?" In all four European countries surveyed (United Kingdom, Germany, Netherlands, and France), wanting to work more was substantially more common than wanting to work less.

In part because of the importance of phrasing, economists are inclined to be skeptical of answers to hypothetical questions such as those used in all of these survey questions on preferred hours. Unfortunately, there is only limited experience in North America with organizations allowing workers to voluntarily reduce work effort in return for a pay reduction. Nevertheless, it does not support the finding that a large fraction of the population would give up income for more vacation. Best (1981) reports that Santa Clara County, California, faced with severe budget cutbacks in 1976, offered workers the option of a 5, 10, or 20 percent pay reduction in return for an increase of 10.5, 21, or 42 days of vacation. We note that, given the existence of holidays and fixed fringe benefits, this is somewhat more favorable to workers than a proportionate reduction in compensation. Seventeen percent of workers increased their vacation. Best also reports that about 16 percent of lawyers in the public defender's office take a three-month sabbatical each year in return for a 25 percent salary reduction. Because interest in less work is more common among higher earning workers in both U.S. and Canadian surveys, these experiences do not suggest a large latent demand for reduced work hours.

Perhaps the most extensive test was in New York state government, which in 1984 adopted a system of voluntary reduction in work schedules, or V-time. This offered full-time employees the opportunity to reduce their work schedules and salaries by 5–30 percent while remaining in their career-path positions. Leave time and pensions were prorated. Subject to their supervisors' approval, employees could reduce their workday or workweek on a regular basis, take time off intermittently, or "bank" time for use at a later date. The official program guidelines did not specify any "acceptable reasons" for requesting V-time, nor did the application even ask for reasons. V-time was not a once-and-forever choice. Workers could request a V-time arrangement to last for as long or short a period as they wished. Many employees were eligible for V-time, including professional, scientific, technical, managerial, and "confidential" employees.

From the perspective of trying to discover a latent demand for reduced working hours, the program could hardly have been more ideal. Its extreme flexibility gave employees themselves the choice of the timing and duration of cutbacks. Nevertheless, very few people actually requested V-time. The number of participants never represented more than 2 percent of employees in the jobs covered by the program. The most common uses of V-time were for temporary maternity and family leaves. As of October 1993, there were only 588 V-time participants, less than 1 percent of the eligible employees.

THEORIES OF HOURS CONSTRAINTS

There is relatively little information on the actual number of hours that individuals wish to work. Based on Kahn and Lang (1995), the average Canadian would like to work 8 percent more hours, comparable in magnitude to the loss in work time due to unemployment. Understanding hours constraints is therefore potentially extremely important. Below we summarize four primary theories in the literature that may explain why workers are constrained to work more or less than they desire.

Long-Term Contracting

Lazear (1979, 1981) has argued that long-term contracts lead to a divergence between the wage and the value of marginal product (VMP). This leads to a conflict between the hours that would be chosen by the worker and firm. Workers will wish to work until the marginal value of leisure equals the wage. Firms will want workers to work until their value of marginal product for the last hour worked equals the wage. Efficiency requires that hours be set so that the marginal value of leisure equals the value of marginal product for the last hour worked. If the value of marginal product from an hour worked is independent of hours worked, it follows that whenever the wage exceeds VMP, workers will be constrained to work less than they want. Conversely, when VMP exceeds the wage, workers will be constrained to work more than they want.

Lazear develops his argument in the context of an agency model. In this model, workers post a bond, in the form of a low starting wage, that is later returned to senior workers in the form of wages that exceed their VMP. Thus, in the agency model, junior workers should be constrained to work more than they wish while senior workers should be constrained to work less than they wish.

In contrast, in many specific-capital models (Becker 1975), workers and firms invest jointly so that junior workers are paid more than their VMP. The firm recoups its investment by paying senior workers less than their VMP. Thus, in the specific-capital model, junior workers are constrained to work less than they wish while senior workers are constrained to work more. Kahn and Lang (1992, 1995) discuss hours constraints in the agency and specific-capital models more fully.

Other long-term contracting models also imply hours constraints. For example, in Harris and Holmstrom (1982), firms and workers are uncertain about how productive the worker will turn out to be. Firms offer insurance contracts in which they promise not to reduce wages. Information about productivity is revealed gradually to the market. Workers who turn out to be unproductive end up being overpaid, while the wages of more productive workers are bid up. As with other long-term contracting models, this can be shown to imply hours constraints. On average, low-seniority workers are paid less than their VMP, because firms are collecting insurance premiums. However, on average high-seniority workers are paid more than their VMP, because firms have stopped collecting insurance premiums and are making insurance payments to low-productivity workers. Consequently, on average more senior workers will be constrained to work less than they wish.[4] Thus, in both the Harris/Holmstrom and agency models, the tendency to want additional hours rises with seniority.

Hedonic Models of the Wage/Hours Locus

For most people, going to work involves substantial fixed costs. Regardless of how long the individual remains at work, she or he incurs the cost of commuting. Once at work, there may also be set-up costs—for example, the time it takes to boot the computer. Therefore, it is no surprise that we observe few workers who are employed for extremely short time periods, because workers would demand a high hourly wage while firms would only be willing to offer a very low one. At the other end of the spectrum, workers who worked very long hours would suffer from fatigue. The workers would require high wages to compensate them for working such long hours, but firms would be unwilling to pay high wages to such workers because their productivity would be low.

More generally, if we were to plot the average hourly wage workers would require to compensate them for different weekly hours of work (i.e., their indifference curves in wage/work-hours space), we would expect the indifference curves to be U-shaped with moderate hours of work requiring less average hourly compensation than very short or very long workweeks. In contrast, if we were to plot the average hourly pay firms would be willing to pay for different weekly hours of work (i.e., their iso-profit curves in wage/hours space), we would expect them to be hump-shaped with moderate hours of work more compatible with higher average hourly wages than either very long or very short workweeks.

If all workers and all firms are identical and there is free entry, equilibrium is at the point of tangency between the indifference curves and the zero-profit iso-profit curve. This point is efficient. Given the options available, no worker or firm wants to change hours.

When workers and firms are heterogeneous, the tangencies of the indifference and iso-profit curves will trace out a hedonic wage/hours locus. Workers who want short hours will be matched with jobs in which short hours are relatively advantageous to the firm. The shape of the wage/hours locus is largely indeterminate. It may be linear, hump-shaped, U-shaped, or wiggly. Regardless of the shape, each firm offers a job with the most profitable wage/hours combination given this locus. Each worker chooses his or her most preferred job given the same wage/hours locus. Again, the equilibrium is efficient, and no worker or firm wants to change hours.

The survey questions described in the previous section typically do not ask workers whether they would prefer to move to a different spot along the wage/hours locus. Instead they ask if workers would like to change hours if they could work at the same hourly rate. Because workers may not have the option of working a different number of hours at the same hourly rate in the hedonic model, they may well prefer to change hours if given this option.

In order to know whether workers will want more or fewer hours at their usual hourly wage, we need to examine the relation between their marginal wage and their average hourly wage. Workers choose to work up to the point at which the marginal wage is equal to their marginal value of leisure. If the average wage exceeds the marginal wage, it will therefore also exceed the marginal value of leisure, and they will desire additional work at that wage. On the other hand, if the marginal wage exceeds the average wage, the average wage will be less than the marginal value of leisure, and workers will prefer to reduce their hours if they can do so at their average hourly wage.

Whenever the average hourly wage is greater than the marginal wage, the hourly wage will be declining with hours worked. Conversely, if the average hourly wage is less than the marginal wage, the hourly wage will increase with hours worked. Therefore, the hedonic model predicts that workers will want more hours if they are on an increasing section of the wage/hours locus and fewer hours if they are on a decreasing section.

Models with Rigid Wages

Hours constraints may arise when workers and firms sign fixed-wage contracts that allow firms to set hours. While fixed-wage contracts are commonly seen empirically, they lack a theoretical foundation. Although insurance-based models would seem the logical theoretical underpinning for fixed-wage contracts, these models suggest that salary, not wages, should be fixed. Similarly, efficiency wage models imply efficiently set wages rather than fixed ones.

Despite its theoretical deficiencies, a fixed-wage model is attractive because it suggests that hours constraints can be viewed as a continuum where hours fall as demand falls, and unemployment is but an extreme. Without a formal theoretical model, it is impossible to make firm statements as to the predictions that follow from this view of hours constraints. Nevertheless, we would expect that in such a model, the desire to work less would be positively correlated and the desire to work more negatively correlated with measures of excess demand in the labor market.

Hours as a Screening Device

Rebitzer and Taylor (1996) develop an explanation of why there might be a shortage of short-hour jobs in certain occupations. The motivation for their model is law associates. Rebitzer and Taylor argue that requiring long hours is a screening device for individuals with low disutility of effort. Because potential partners care about being part of a firm with hard-working partners and because partner effort is difficult to monitor, law firms benefit from requiring that associates work long hours. Provided that disutility of effort and disutility of hours on the job are correlated, reducing hours may create an adverse selection problem by attracting less hard-working individuals to the firm.

EVIDENCE FOR AND AGAINST THE THEORIES

Long-Term Contracts

Kahn and Lang (1992, 1995) report that in both the PSID and the Canadian SWR, wanting to work additional hours is negatively related to seniority. Table 5 presents some representative results from the Canadian data. Almost half the most junior workers want *more* work, compared with roughly 20 percent of the most senior workers. In contrast, only about 10 percent of the most junior workers but twice as many senior workers want to work less. The relation between seniority and the constraints favors models such as firm-specific capital, in which wages grow less rapidly than VMP. On the other hand, at no seniority level does the average worker want less work or does the number of workers wanting less work exceed the number wanting more. This suggests that long-term contracting cannot be the sole explanation for hours constraints. If it were, the results would imply that wages exceed VMP at all seniority levels which is inconsistent with profit maximization.

The Hedonic Model

The distribution of actual hours appears to be responsive to desired hours, suggesting that matching takes place in the labor market as predicted by a hedonic model. Kahn and Lang (1995) report that in Canada, usual hours worked increase by half an hour for every hour increase in desired hours. This is true both for individuals and for mean usual and desired hours across provinces. In the European Union survey, among 11 countries, the fraction of workers wanting to work more than 45 hours is correlated with the fraction actually working more than 45 hours ($r = 0.42$) and even more so if we exclude Portugal ($r = 0.80$). The correlation between the fraction wanting to work less than 20 hours and those working less than 20 hours is even greater ($r = 0.94$, including Portugal).

On the other hand, the matching seems to be only imperfect inasmuch as it improves over time. In the Canadian results reported in the first part of Table 5, the fraction of workers who do not want to change their work hours rises from 43 percent among the lowest tenure group

Table 5 Proportion of Workers Experiencing Binding Hours Constraints by Job Tenure and Short/Overtime (Canada)[a]

Job tenure (months)	More work	Less work	Satisfied with hours	*N*	Number of hours[b]
1–3	47.9	9.2	42.9	740	5.45
4–6	46.5	11.3	42.2	360	5.30
7–9	41.5	12.2	46.3	200	4.66
10–12	41.9	17.9	40.2	259	4.00
13–24	36.6	15.9	47.4	660	3.24
25–36	35.8	18.0	46.5	435	3.28
37–48	37.5	16.9	45.5	446	3.18
49–60	35.0	18.9	46.0	472	2.98
61–120	34.9	17.7	47.4	1598	2.80
121–240	25.9	20.8	53.3	1500	1.60
> 240	22.0	20.2	57.9	574	1.21
Short/overtime					
On short-time	37.4	27.0	35.6	97	4.60
Normal hours	33.4	17.3	49.3	6167	2.88
More than usual	39.1	16.9	44.1	980	3.17
All	34.2	17.3	48.5	7244	2.94

SOURCE: Kahn and Lang (1995).

[a] Based on the Survey of Work Reduction supplement to the Canadian Labour Force Survey, June 1986.

[b] "Number of hours" is the average number of additional hours desired by members of the group. All observations are weighted by their sampling weight.

to 58 percent among the highest tenure group. This suggests that either workers adjust their tastes over time or that dissatisfied workers leave for jobs with hours requirements that conform better to their tastes. Using the PSID, Altonji and Paxson (1988) find that U.S. workers' responses regarding desired hours help predict whether workers will subsequently shift to longer or shorter hour jobs.

Additional evidence suggesting only imperfect hedonic matching is found in the fact that substantial fractions of part-time workers would prefer full-time work, and vice versa. One likely interpretation of these statements is that these people would prefer to be at a different point along the wage/hours locus but cannot. Thus, although the hedonic model may give insight into how hours and wages are determined, the empirical evidence suggests that there is substantial mismatching.

To test the hedonic model more formally, Kahn and Lang (1996) use the June 1986 Canadian SWR to estimate the wage/hours locus and an hours-constraints equation simultaneously. We test whether the pattern of hours constraints conforms to the hours constraints that should be generated by the wage/hours locus under the hedonic model. A pure hedonic model is easily rejected. We do not find that workers wanting more hours are in the downward-sloping part of the wage/hours locus and those wanting fewer are in the upward-sloping part.

However, when hours constraints are allowed to depend on seniority (as in the firm-specific model) as well as on the slope of the wage/hours locus (as in the hedonic model), the empirical model fits surprisingly well. Kahn and Lang (1996) plot the actual wage/hours locus estimated from income and hours data and the wage/hours locus predicted on the basis of hours constraints and seniority. The curves are quite similar except in the region beyond 60 hours, and the poor fit in this region is due to the fact that average weekly earnings actually fall beyond 58 hours. Excluding these long-hour workers, who tend to be low-tenure workers in managerial or administrative positions who are presumably investing in their career, the equality of the two equations cannot be rejected.

Thus, we cannot reject a model of hedonic matching combined with long-term contracting due to firm-specific skills.[5] This is somewhat surprising in light of the evidence of imperfect matching cited above. It is also difficult to reconcile the sharp spike in the distribution of hours with the pure hedonic model. To some extent, the failure to reject the hedonic model must reflect relatively inefficient statistical techniques forced on us by the imperfect data.

Models with Rigid Wages

Hours constraints are correlated only weakly if at all with measures of labor demand. Because the Canadian data are cross-sectional, we measure variation in labor demand two ways. First, we utilize the fact that workers who respond that they are on short-time (reduced hours) will tend to be in firms and/or industries experiencing unusually low demand, while workers who are working more than their usual hours will tend to be in firms and/or industries experiencing unusually high demand. The bottom panel of Table 5 shows the hours constraints for 1) workers working less than usual because of short-time,[6] 2) workers working their normal hours, and 3) workers working more than their usual hours. The desire to work fewer hours is most common among workers on short-time. Similarly, the desire to work more hours is most common among those working more than their usual hours. The results suggest that when the establishment faces low demand, workers who want to work fewer hours take advantage of the situation to reduce their hours. Similarly, when the establishment needs additional hours, workers who want more hours are able to increase their hours. Thus, hours constraints cannot be interpreted as cyclical underemployment being imposed on unwilling workers.

We also capture labor demand from the cross-sectional Canadian data through regional unemployment rates. Here, too, the relation is weak. Among Canadian provinces, there is a positive relation between average additional hours of work desired and the unemployment rate, but it falls well short of conventional significance levels (Kahn and Lang 1995). Similarly, in the U.S. PSID data, controlling for other factors, the local unemployment rate is positively related to the desire for more work, but the coefficient is generally insignificant (Kahn and Lang 1992).

Using time-series data from the PSID in Table 2, there is some evidence of a relation between the prevalence of hours constraints and the national unemployment rate. The proportions of both men and women wanting to work more are each positively correlated with the civilian unemployment rate for men age 20 and over, although this result is not robust to including a time trend in the case of women. On the other hand, the proportion wanting to work less is not significantly related to the unemployment rate, and the correlation is positive.

Hours constraints, however, are related to recent personal unemployment experience in both Canada and the United States, even controlling for experience and seniority (Kahn and Lang 1992, 1995). One explanation for this result is that workers who obtained their jobs after an unemployment spell are less well matched than those who "chose" new jobs and experienced no intervening unemployment. A second explanation could be that people tend to want long hours after an unemployment spell because they have run down their assets, but that mismatching makes it unlikely that they actually obtain these long hours.

Hours as a Screening Device

The Rebitzer and Taylor (1996) screening model predicts only overemployment, rather than both overemployment and underemployment. It thus cannot explain the desire to work additional work hours that is so common in the United States and Canada. Rebitzer and Taylor found their theory to have explanatory power for lawyers. It may apply as well to other similar occupations. In both Canada and the United States, people who desire to work fewer hours tend to be higher-earning workers, i.e., more educated, in more skilled occupations, etc. (Kahn and Lang 1991; Drolet and Morissette 1997). These may indeed be the occupations where hours are a signal of effort.

HOURS CONSTRAINTS, THE DISTRIBUTION OF HOURS, AND UNEMPLOYMENT

In sum, it appears to us that hours constraints are best understood in the context of a matching model in which wages do depend on hours as in hedonic models, but in which the matching is imperfect. (There may also be long-term contracting.) An imperfect matching model would also allow us to evaluate hours policies in the context of a model in which unemployment as well as vacancies can arise. In this section, we take some tentative steps toward analyzing the impact of mandated hours reductions in the context of such a model. The model we use is a simple extension of the Butters (1977) equilibrium search model. We describe it only informally.

In labor market variants of the model (Hosios 1986; Lang 1991), each firm decides simultaneously whether or not to make an offer to a worker and, if so, what wage to offer. Making an offer entails paying a fixed cost. Under certain circumstances this may be interpreted as the cost of renting a machine prior to trying to hire a worker. The worker chooses the firm that offers him or her the highest wage. Because some workers may get only one offer and others may get multiple offers, the equilibrium involves a wage distribution. Each firm recognizes that offering a higher wage raises the probability of hiring the worker but lowers profits conditional on getting the worker. Because some workers randomly fail to receive any offers, there is unemployment. Similarly, some firms' offers are turned down, resulting in vacancies.

To take account of hours constraints, we extend the model in a simple way. First, we assume that firms make tied wage/hours offers. The worker chooses the firm offering the highest utility level provided that utility level exceeds some reservation utility level. For simplicity, we assume that the value of marginal product per hour v is independent of hours worked and that the utility function is given by $u = \log(wh) + (T - h)/\beta$ where wh is (labor) income, h is hours worked and $(T - h)$ is leisure. This utility function has the property that desired work hours equal β and are independent of the wage rate.

We note that the resulting equilibrium is very much a theoretical counterpart to Dickens and Lundberg's (1993) study of constrained labor supply in that workers choose from a limited and stochastic number of wage/hours offers. In contrast with that paper, we allow for unemployment.[6]

The firm chooses w and h to maximize expected profits which are given by

$$\text{(1)}\quad E(\pi) = P(u)\,(v - w)\,h - d,$$

where u is the utility associated with the offer, P is the probability of the offer being accepted, and d is the fixed cost of making an offer. The equilibrium is characterized by a distribution of wages and hours, and of utilities with corresponding values of P.

It is relatively straightforward to prove the following:[7]

1) All firms offer hours in excess of β. In other words, all workers would respond that they would want to work less at their usual hourly wage.

2) There is a distribution of hours and wages. The hourly wage is monotonically declining in hours. In fact, wh is constant.

If we were to observe the facts 1 and 2 without the perspective of the model, the logic for mandating hours restrictions would seem compelling. The atheoretical perspective would be as follows. Workers want shorter hours. If hours could be reduced but salaries maintained (i.e., *wh* constant), this even suggests that firms are equally happy with both situations. Moreover, the quantity of labor demanded would rise, thereby reducing unemployment. Thus mandating a shorter workweek to reduce unemployment would appear to be a "sure-fire winner."

Unfortunately, within the context of the model, that policy assessment turns out to be completely wrong. Again, it can be shown [see Appendix] that:

1) mandating lower hours increases unemployment,
2) mandating lower hours decreases wages, and
3) mandating lower hours is welfare deteriorating in the sense of Pareto.

Thus, in contrast to the conclusion we might be tempted to draw, mandated hours restrictions will not be desirable. The Appendix works out a numerical example that illustrates these results.

Our choice of utility function and production function were designed to generate an equilibrium in which workers express a desire to work fewer hours. We chose this example because we believe that this equilibrium would appear to provide a strong *a priori* case for mandated hours restrictions when examined atheoretically. It is easy to choose utility and production functions such that workers desire to work more hours at their usual hourly wage.

At the cost of some complexity, Lang and Majumdar (2000) extend this model to allow for heterogeneous preferences and thus for imperfect matching. Because each worker chooses from only a limited number of jobs, matching is imperfect. Nevertheless, workers preferring low-hours jobs tend to end up in jobs with low hours since they take these jobs whenever a choice is available. They find that hours restrictions can increase or decrease unemployment. The principal welfare effect is distributional. Workers who prefer jobs with short hours are better off while those who prefer longer hours are worse off.

IMPLICATIONS FOR MANDATED HOURS RESTRICTIONS

Neither the empirical nor the theoretical case for mandating hours restrictions to increase work sharing and reduce unemployment is compelling. In the United States and Canada, there is very little evidence that workers are interested in accepting less pay in return for more leisure. The situation in Europe may be different.

In the introduction to this chapter, we argued that we cannot evaluate theoretically a policy designed to lower unemployment within the context of a model which assumes away unemployment. Neither the long-term contracting nor the hedonic matching model predicts any unemployment, while "models" of rigid wages have no theoretical underpinnings with which to evaluate policy. However, the imperfect matching model sketched in the previous section predicts both hours constraints and unemployment. We have established that such models do not justify casual support for mandated hours restrictions. Indeed, our simple model suggests that they may be welfare-deteriorating and lead to wage losses and even more unemployment. However, mandated hours restrictions may be welfare-improving in some situations such as law firms where hours serve as a screening device.

Moreover, any attempt to legislate a reduced workweek and promote work-sharing will undoubtedly increase the pervasiveness of dual job-holding, at least within the United States and Canada. Whatever theoretical model is assumed, there is also likely to be a change in the wages paid for jobs with different levels of required hours. Any analysis of the effect of mandated hours restrictions must take these effects into account.

Notes

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1. We exclude proprietary surveys for which basic information on question wording and sample design are not available and surveys with questions that do not explicitly suggest an earnings/hours trade-off. A 1993 Gallup Poll asked workers their actual and desired hours. Mean actual hours were reported as 42.5 while desired

hours were 36.7. However, 16 percent of workers responded that they preferred zero hours, suggesting that these respondents were not thinking of an hours/salary trade-off. Excluding this 16 percent of respondents, mean desired hours was 43.7.

2. The exact questions asked were
    - "In the next two years, would you take a cut in pay if you received more time off in return?" Follow-up if no: "Why not?" Follow-up if yes: "What percent of your pay would you give up to have more time off?" Accompanying these questions was a table and examples to help the respondent think about how much money an *x* percent pay cut represents, and how much time an *x* percent hours cut represents.
    - "Another way to gain more time off is to trade all or some part of your pay increase. Would you trade some of your increase in the next two years for more time off? For example, gain 5 percent more time off instead of a 5 percent pay raise?" Follow-up if yes: "How much of your increase in the next two years would you take as time off?"

    The questionnaire proceeded with a set of questions on how the person would prefer to reduce work time (e.g., fewer hours per day) and about reasons that person preferred to work less. The questionnaire then continued with:

    - "If you continue to be paid at the same rate of pay that you are now, would you work more hours for more pay?" Follow-up if yes: "How many more hours per week would you want to work?"
3. Many people gave inconsistent answers. For instance, almost one-quarter of those who said they were willing to take a pay cut for fewer hours also said they were not willing to forego pay increases for fewer hours, and half of these actually said they would like to work more hours for more pay.
4. Strictly speaking, only workers in the final work period have stopped paying insurance premiums and only in this period are high productivity workers paid exactly their marginal product. In other years, they are underpaid because of insurance premiums.

    Allowing mobility reinforces the tendency for more senior workers on average to be paid more than their marginal product and hence to prefer more work hours. Workers revealed to be high productivity will be indifferent among all firms while workers revealed to be overpaid will prefer employment at their present employer. Consequently, highly productive workers are more likely to change jobs than are overpaid workers, further adding to the average overpayment of senior workers.
5. The fact noted earlier that seniority decreases the desire to work more hours is also consistent with the firm-specific capital model.
6. Dickens and Lundberg (1993) is primarily an empirical paper, but it incorporates a structural model. Because their data set included only employed people, they did not model unemployment.
7. Differentiate Eq. 1 with respect to both w and h. Dividing one first-order condition by the other and rearranging terms gives $w = \beta v/h$. Since firms can only make profits if $w < v$, this requires that $h > \beta$. Moreover, it implies that $wh = \beta v$, a constant.

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Appendix

Consider a worker who obtained utility U_0 in the unconstrained equilibrium. The firm chooses the profit-maximizing combination of wage/hours for that U_0. Let $\pi_a = (v - w)h - d$ be the profit obtained conditional on hiring a worker achieving U_0. (Recall that the probability of the offer being accepted P is dependent only on U.) When hours are constrained, if the firm were still to hire a worker who received U_0, the new profit must be lower than the unconstrained profit.

The zero profit condition requires that $P(v - w) = d$. Therefore, when hours are constrained, $\pi_c \leq \pi_a$ implies that $P_c(U_0) \geq P_a(U_0)$. For this to be true, the likelihood of a worker obtaining utility greater than U_0 must be lower (or equal) when hours are constrained than when they are not.

Hence, workers are worse off (or, more formally, no better off) in the constrained equilibrium. This argument applies to all utility levels, including the reservation utility, the lowest utility offered. If the likelihood that a worker who is offered the reservation utility accepts the offer is greater in the constrained solution, then the likelihood of a worker receiving a utility greater than this minimum reservation level is lower; so we will see higher unemployment rates.

It may also be helpful to work through a numerical example. Suppose that β equals 40 (so that workers' desired hours equal 40 as well), v equals 10, and d equals 20. We set the reservation utility so that in equilibrium the maximum number of hours in any offer is 60. The following can be derived: All workers are offered an income of 400. Hours offers range from 42 to 60; wage offers range from 6.67 to 9.52. The unemployment rate is 10 percent. When hours are set exogenously at 42, wage offers range from about 6.07 to about 9.52. The unemployment rate is about 12 percent. In a standard competitive model with this same utility function and $v = 10$, a profit-maximizing firm sets hours $= 42$. Imposing the "competitive solution" dramatically lowers employment and wages.

Part IV

9

Measuring the Effects of Short-Time Compensation on Workforce Dynamics

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This chapter examines existing research on the relationship between short-time compensation (STC) use and patterns of labor market adjustment by firms. We note that the research often fails to model the actual choices available to firms and their workers in which both reductions in compensated hours and layoffs are feasible. We show that the strategies used previously to calculate the extent to which time on STC affects time on layoff have yielded a wide range of estimates. We also explore new data from a recently completed evaluation of STC programs in the United States. These data confirm the need to study labor force adjustment strategies along multiple dimensions, but they also illustrate the difficulties involved in assessing the precise impact of STC use on workforce adjustment.

North American employers—particularly those in the United States—have often been alleged to prefer layoffs to hours reductions when responding to output or cost shocks (Burdett and Wright 1989; Feldstein 1976, 1978). Feldstein, for example, highlighted two features of the U.S. system that encouraged firms to opt for layoffs during cyclical downturns: 1) unemployment insurance (UI) benefits were, at that time, nontaxable to workers who received them, and 2) the UI payroll tax was not fully "experience rated" so that firms did not incur the full costs of benefits paid to their workers. According to Feldstein, these features created a strong incentive for firms to use layoffs to reduce the workforce. The author's empirical estimates, together with subsequent estimates by others (for example, Topel 1983), tended to support the notion that layoffs, especially temporary ones, were correlated with the level of the "layoff subsidy" the UI system provided.

In an effort to identify other ways in which the U.S. system of unemployment compensation may encourage layoffs, several other authors have focused on the way the UI systems treat hours reductions. In many European countries, workers who are placed on reduced hours are eligible for a prorated share of their unemployment benefits. Until recently, however, the availability of such STC was severely limited in both the United States and Canada.

In this chapter, we critically examine the methods that have been used to assess the effects of STC on total workforce reductions. We begin by briefly discussing the existing theoretical approaches to this topic. We then describe a number of previous attempts to develop empirical evidence on conversions between STC and layoffs. We devote considerable attention to evaluating methodologies that have been used for this purpose, because they have yielded widely differing results. In the next section, we present some new empirical evidence on the relationship between STC use and layoffs from our recently completed evaluation of STC programs in the United States. This research further confirms the importance of this relationship and illustrates the difficulties in measuring it precisely. Finally, in the last section, we offer some general conclusions about our state of knowledge on the ability of STC-type programs to influence firms' workforce adjustment patterns.

THEORETICAL MODELING

To understand how employers' and workers' preferences interact when labor input is reduced during periods of changing demand, the development of an employment contract model is required. The approach usually taken in the literature draws on the early implicit contracts model developed by Azariadis (1975). This model views workers and employers as engaged in a bilateral bargaining process. An efficient outcome from the process is a set of choices that maximizes each party's well-being (that is, profits for firms and utility for workers), given the choice of the other party. In its most general form, this model predicts that risk-averse workers will generally prefer hours reductions to layoffs during economic downturns. This preference can be altered by technical aspects of a firm's production and cost functions. For example, if the firm's adjustment costs are asymmetric with respect to changes in hours and employment, different patterns may be optimal. High fringe benefit costs, especially those that are "quasi-fixed," may deter downward adjustments in hours. On the other hand, hiring costs, such as those related to the search for workers or to training and acquiring job-specific human capital, may deter layoffs. Choices may also be affected by imperfect substitutability between hours and employment in the production process.

A few theoretical papers (Wright and Hotchkiss 1988; Burdett and Wright 1989; Jehle and Lieberman 1992) have attempted to explore how the availability of various UI and STC options may affect hours–employment choices. Under a stylized "American" system, UI benefits are assumed to be payable only if the worker is fully separated from the firm. Alternatively, a stylized "European" system of compensation is assumed to provide benefits only for reductions in normal working hours. Models adopting this approach suggest, not surprisingly, that the American system encourages firms to opt for layoffs during downturns in demand, whereas firms operating under a European system, *ceteris paribus*, favor shortened workweeks. In both cases, the work reduction incentives derive primarily from the incomplete experience rating of UI benefit payments. Hence, some authors favor a move toward more complete experience rating as the primary way to ameliorate inefficient labor input choices encouraged by UI benefits (see, for example, Burdett and Wright 1989).

The all-or-nothing nature of the stylized UI systems in the theoretical literature makes it difficult to apply these models directly to actual data. In Europe, the bulk of UI benefits goes to workers who are fully laid off rather than on reduced hours, and in North America, Canada has a national STC program, as do 18 U.S. states.[1] Furthermore, most examinations of firms' adjustment patterns on the microeconomic level, especially in North America, have found that individual firms use both layoffs and work-week reductions to reduce labor utilization (Kerachsky et al. 1986; Employment and Immigration Canada [EIC]1993). Hence, a clear first step in the development of theoretical models that can be estimated from real world data is to generalize the nature of adjustment options faced by the firm.

Adapting these theoretical models to allow UI eligibility for both hours reductions and layoffs seems straightforward, although the published literature has not attempted this adaptation in its full generality. Presumably, such a generalized model would predict that both types of reductions would be encouraged by the availability of benefits, with firms choosing the type of reduction or combination of reductions on the basis of their own cost and productivity considerations.

Differences in the generosity of benefits available under the two unemployment options could also affect observed workforce adjustments, a point made forcefully in the recent paper by Van Audenrode (1994). Indeed, the increased flexibility of a system that compensates both hours reductions and layoffs might encourage additional compensated unemployment, relative to a system that provides UI eligibility for only one type of workforce reduction.

Aggregate studies of U.S. labor market dynamics over the business cycle provide evidence on this issue. These studies suggest that labor hoarding may have accounted for between 4 and 9 percent of total employment during downturns in demand, probably because of high adjustment costs associated with layoffs (Hamermesh 1993, p. 185). Some portion of this excess labor would probably find compensated hours reductions attractive after STC becomes available.[2] The extent of these incentives depends on variations in fringe benefit costs available from hours reductions and the degree to which STC benefits are effectively experience rated, among other things. Much of the existing empirical research on STC has not, however, addressed the increasing flexibility suggested by theoretical predictions.

EMPIRICAL STUDIES MEASURING THE EFFECT OF STC USE ON WORKFORCE DYNAMICS

Because one of the primary goals of STC is to reduce the number of laid-off workers, most previous researchers have tried to estimate differences in the dynamics of workforce reductions that result from STC usage. Although many different approaches have been taken to analyzing such reductions, we summarize these approaches in what we call the "layoff conversion rate." This measure reflects the degree to which unemployment compensation under an STC-type program substitutes for unemployment compensation under a regular UI program.[3] A conversion rate of 1.0 (a value frequently assumed in the literature) implies perfect substitution: each hour of STC substitutes for precisely an hour of layoff. Layoff conversion rates greater than 1.0 imply that the compensated unemployment from layoffs avoided because of STC exceeds the compensated unemployment from STC itself. Conversion rates less than 1.0 imply that firms had greater total workforce reductions with STC than they would have if STC had been unavailable.

In reviewing previous research, we focus narrowly on ways in which conversion rates have been treated. Many of the studies we discuss here have contributed significantly to an understanding of the other important issues such as 1) the effects of STC on employer and employee satisfaction, achievement of affirmative action goals, and worker productivity; 2) the effects of program legislation and administration on STC participation rates; 3) the relative costs and benefits of STC usage to employees, firms, the UI trust fund, and society; 4) seasonal, cyclical, and repeat use of STC; and 5) the use of STC by firms undergoing structural change. We do not summarize research on these issues here, however.[4] Rather, we concentrate solely on the methodology researchers have used to examine layoff conversion rates. Hence, we are not attempting to explore the full social costs and benefits of STC compared to layoffs. For many issues involving STC desirability, however, estimates of the layoff conversion rate play an important, even central, role. For example, measuring any potential social benefit from STC use requires some way of estimating how many layoffs would have occurred in the absence of the program so that the analysis can be conducted on a "per layoff equivalent" basis. Much of the prior research on STC has not been especially careful in adopting such a consistent basis. Our research focus on measurement of workforce dynamics therefore serves to highlight a primary source of the differences in conclusions from previous research.

Experimental studies to determine the layoff conversion rate have not been feasible, so researchers have typically used one of three nonexperimental approaches to estimate the workforce reduction that would have occurred if firms had not used STC: self-reporting, the explicit or implicit assumption that the conversion rate is precisely 1.0, and estimation using matched samples of firms that did and did not use STC during some period. In the next three subsections, we discuss these approaches and some of the research based on them. For ease of comparison among the studies, those that we shall discuss are briefly summarized in Table 1.

Studies Based on Self-Reporting

One way of estimating the number of layoffs avoided because of STC usage is based on firms' self-reports. This method uses two main data sources: firms' plan applications and surveys of employers. When a firm applies for an STC plan approval, it typically has to specify the

Table 1 Studies of the Effects of Short-Time Compensation on Workforce Dynamics

Study	Data	Method	Conversion rate
New York Department of Labor (1994)	New York STC Firms	Self-reporting	>1
Best (1988)	Canadian and California STC firms, administrative data, and simulations	Self-reporting	1.0
Vroman (1992)	Germany, 1970–1991, administrative data	Assumed	1.0
Abraham and Houseman (1994)	France, aggregate manufacturing data	Assumed	1.0
Van Audenrode (1994)	Five European countries, aggregate employment data	NA[a]	NA
Kerachsky et al. (1986)	Arizona, California, and Oregon	Comparison matching	0 to 1.0
Employment and Immigration Canada (1993)	Canada, worksharing and comparison firms in 1989 and 1990	Comparison matching, self-reporting	>1.0
Berkeley Planning Associates and Mathematica Policy Research, Inc. (1997)	Five states	Comparison matching	See text

[a] NA = not applicable.

number of layoffs that its use of the plan would avoid. Because a firm needs administrative approval for its STC plan to become effective, it may have an incentive to overstate the number of layoffs that would occur if its STC plan is not approved. The validity of this counterfactual cannot be tested directly.[5] EIC (1993, pp. 58–59) found that firms typically overstate their planned workforce reductions; that is, the actual reduction is less than the planned reduction. This study also found that a significant number of post-STC layoffs occurred. These factors suggest that using firms' self-reported statements on the number of layoffs that will be averted if an STC plan is approved overestimates the effects of STC on layoffs. EIC (1984, pp. 116–118) found that retrospective interview responses on layoffs that would have occurred had STC not been available were 23 percent lower than firms' self-reports of planned layoffs from STC plan applications.

A second source of self-reported information on averted layoffs is survey data from firms that have used STC, such as those reported in EIC (1984). These data may suffer from the same problem as self-reported data on STC plan application forms—firms are asked to hypothesize about how many employees they would have laid off had STC been unavailable. A firm might never have laid off any employees (as the "labor-hoarding" literature suggests), and it might also have laid off significantly fewer or more full-time employees than the full-time equivalent (FTE) value of the STC reduction implemented.

Most states with STC legislation have not conducted explicit studies to estimate the number of layoffs averted by STC usage. Their research interest has focused on other aspects of the STC adoption and financing processes. The New York Department of Labor, however, has estimated the number of layoffs averted in each year since 1988. In 1994, for example, 445 New York firms had a total of 9,284 employees on STC plans, which paid out $3.6 million in benefits. These firms reported that, because of STC use, almost 4,000 layoffs were averted. Using average benefit levels and unemployment durations for laid-off workers, the state estimated that about $10.8 million was saved in UI benefits in 1994 alone. If we assume perfect experience rating, these results suggest that, on average, firms saved three dollars in potential UI taxes for every dollar paid out in taxes to support STC benefits. At face value, this calculation seems implausible and is inconsistent with the low observed utilization rates for STC in New York and elsewhere.

As we shall see, the result also conflicts with most other empirical evidence on STC.

Studies Assuming a Conversion Rate of 1.0

A second way to estimate the number of layoffs avoided because of STC usage is based on calculating the FTE workforce reduction directly from administrative records (or survey data) on the number of employees collecting STC and their workweek percentage reductions using an assumed conversion rate of 1.0. For example, if a firm has 10 employees on a 40 percent workweek reduction for five weeks, the researcher assumes that the STC plan averted four layoffs, each of a five-week duration.

It is important to recognize that several assumptions are inherent in the "one-for-one" conversion rate assumed in this type of estimation. Most important, the calculation assumes a linearity that often does not exist in either production technologies or employment policies. In standard economic theory, the firm maximizes profits by adjusting labor and other inputs. Because STC may change many factors in the profit function—such as productivity, labor costs, and logistical constraints—it is unlikely that firms would choose the same person-hours of labor input under both shortened workweeks and full-time layoffs. Assuming a one-for-one conversion rate suggests that firms are not responsive to the theoretical advantages and disadvantages the researcher is trying to estimate (or that the advantages and disadvantages cancel each other out). In many situations, however, researchers assume a one-for-one conversion rate primarily because data limitations prevent estimation of the rate directly.

For example, Best (1988) uses a variety of data at the firm level to conduct a comprehensive and innovative evaluation of California's STC program and the old Canadian program. This study presents a good discussion of the factors that can affect the layoff conversion rate, such as firms' ability to resume production more quickly after a downturn if STC is used, laid-off workers' tendencies to leave unemployment for new jobs, and workers' tendencies to oppose STC less than layoffs. Because he had no administrative data on nonparticipating employers, Best relied heavily on employer and employee survey data and simulations to derive his estimates. Although he presented information on employers' and employees' perceptions of the work loss from STC relative to work loss under layoffs, he suggested that these data may be invalid because of inconsistencies between the perceptions of the two groups within firms (p. 76). He concluded that using a one-for-one conversion rate is the most reasonable approach because the data on actual conversion rates are mixed, and these rates may vary significantly over time and by other factors unique to an individual firm.

Best also simulated different estimates of the cost of STC relative to the cost of layoffs for the UI system (including both benefits paid and administrative costs). These estimates, ranging from 1.2 to 3.7, depend on the duration of work-sharing plans, the magnitude of the workweek reduction, and the percentage of STC participants laid off after STC.[6] Best acknowledged that most of the scenarios he presents are uncommon, and it appears that the simulation closest to the average workweek reduction, average duration of the reduction, and average post-STC layoffs provides a ratio of the cost of STC to the costs of equivalent layoffs for the UI system of around 1.6. Because this estimate is based on an underlying assumed conversion rate of 1.0, the extra costs from STC arise from such factors as the higher weekly UI benefits of STC recipients, differential treatment of the waiting week under the two programs, and the additional administrative costs of STC. The result shows that there can be a considerable difference between conversion rate estimates based on equivalent hours of layoff and estimates based on costs to the UI system.

Studies based on aggregate data have tended to use an assumed conversion rate of 1.0 when attempting to estimate the impact of STC usage.[7] This is especially true for studies that have sought to evaluate STC in the European context. For example, Vroman (1992) used administrative data on STC usage in Germany to estimate what employment would have been in the absence of the program during the period 1970 to 1991. To make that calculation, he simply subtracted "full-time equivalent layoffs" experienced by workers on STC from actual employment data, thereby implicitly assuming a conversion rate of 1.0. He found that the cyclical behavior of his adjusted employment series has a closer relationship to the cyclical behavior of U.S. employment than the unadjusted series. Hence, Vroman concluded that the greater availability of STC is an important reason for observed European/American differences. The author presents no empirical evidence to support his conversion rate assumption.

The study by Abraham and Houseman (1994) examined the ways in which job security regulations in Belgium, France, and Germany affect labor force adjustments in response to output shocks. The authors devoted some discussion to the possible influence of STC programs on this adjustment process, although this was not a primary focus of their research. Using aggregate data on hours worked and hours on STC, they show that compensated short-time hours play an important role in hours adjustments in Belgium and Germany. They did not explicitly consider the layoff conversion issue for these countries. Because aggregate data on hours worked were not available for France, however, the authors studied only aggregate employment trends. For France, they constructed a hypothetical employment series "assuming that layoffs were used in lieu of short time." This construction required the authors to assume that layoffs and short-time could be substituted on a one-for-one basis. In common with most of the other literature on European STC programs that we reviewed, the authors offer no empirical support for this assumption. Hence, the issue of precisely how widespread STC availability in Europe affects use of layoffs remains open.

Studies Using Matching Methods

A third way of estimating the numbers of layoffs avoided because of STC use is based on pairing firms that used STC with firms that did not. Difference-in-differences analysis is used to compare the FTE workforce reductions of the STC firms and non-STC firms over time. The critical assumption in this approach is that non-STC firms do not differ systematically from STC firms; that is, unobserved differences between the two groups are independent of treatment status.[8]

A growing set of economic literature has evaluated such nonexperimental evaluation (matching) methods (see, for example, Friedlander and Robins 1995; LaLonde 1986; and Fraker and Maynard 1987). Selecting the pool of potential matches on the basis of similarities in time, geographic area, and observation-specific characteristics is one of the most difficult aspects of matching, and the appropriateness of various criteria for restricting the pool has been debated.[9] For example, researchers may consider limiting the pool of potential comparison firms to those with certain observed levels of compensated unemployment, even though firms that used STC may have chosen not to lay off workers. Thus, researchers who use matching procedures to generate comparison samples may have to make many decisions about what constitutes a good match, without being able to draw on much economic theory as a guideline.

Matching procedures have several practical limitations. First, although matching attempts to control firm-specific differences at the outset of the research design, the variables used for matching may not adequately represent all factors affecting workforce adjustment strategies. The financial health of firms, their labor/management relations, the demand for their products, and their production technologies, as well as trends in these factors, may affect whether firms consider workforce reductions.[10] Data on these factors, however, are extremely hard to obtain; most likely, the variables firms use to make their production (and labor input demand) decisions are known only to the firms themselves. Second, the treatment variable in matching studies must be defined carefully. Because firms change their STC status over time, and enrollment can begin at any time, construction of this variable requires focusing on a particular period (the study period) during which the firm "uses" the program. But that definition must invariably involve some ambiguity when intensity of usage varies. Finally, because comparison firms are chosen to be as similar as possible to STC firms, they are also likely to have participated in the STC program at times outside the study period. Such prior participation may bias the estimated treatment effects.

A 1986 study conducted by Mathematica Policy Research, Inc., (MPR) matched STC firms in Arizona, California, and Oregon to non-STC comparison firms on the basis of their size, UI tax rates, and Standard Industrial Classification (SIC) code (Kerachsky et al. 1986). The comparison firms in each state were chosen from among all firms in the state not using STC during the study period. Empirical estimates from this study found widely varying layoff conversion rates across the three states: California's STC program did not appear to avert any layoffs; Arizona's STC program averted some layoffs, although total unemployment increased for firms using the program; and Oregon's program appeared to have a layoff conversion rate that approximated the 1.0 value assumed in many studies.

Although a matching process was used, STC firms seemed to be in greater economic distress because they had somewhat higher levels of pre-STC compensated unemployment. Although the researchers controlled for this factor in most of their analysis, concerns about their ability to control for pre-STC differences in layoff propensities between the STC and comparison groups resulted in some criticism of their findings. These criticisms focused especially on the finding of no STC impact on layoffs in California and on the possibility that the state of California extracted the data incorrectly (Best 1988, pp. 75–76). Still, the MPR study helped to emphasize the importance of measuring rather than assuming the extent of layoff conversions that STC provides.

The recent Canadian evaluation used a very different methodological approach to matching (Employment and Immigration Canada 1993). In constructing the comparison group, the researchers chose a random sample from administrative records of employees who had been laid off in 1989 or 1990. The firms from which the employees were laid off were screened to ensure that the comparison firms chosen had been in existence for at least two years and that they had "considered" laying off 20 percent or more of the full-time employees in a business unit for nonseasonal reasons.[11] Eligible comparison firms could not have used the STC program in the past, and laid-off employees had to have been recalled within 26 weeks. The analysis included a total of 1,080 firms.

Because firms in the STC and comparison samples were not matched according to specific characteristics (except the screening requirement that comparison firms had to have considered laying off employees), the two samples differed markedly along many of the dimensions that the earlier MPR study used for matching, such as geographic location and industry. Most notably, STC firms were only about one-third the size of comparison firms, on average. Possibly because of these and other differences, many of the results from the analysis of raw data were not supported when regression adjustment techniques were used. In contrast to the MPR study, STC firms in Canada appeared to have been in less economic distress than were firms in the comparison group because employees in these firms had significantly lower pre-STC compensated unemployment than comparison firms' employees.

Despite using a comparison group for some parts of the analyses, the Canadian evaluation did not directly compare layoffs or total compensated unemployment for STC and comparison firms.[12] To determine the number of layoffs averted because of STC, the ratio of STC participants to plan-reported hypothetical layoffs was calculated for 1989 and 1990. Because the firms included had 177,800 employees using STC during these years, and the firms reported that STC averted 67,500 layoffs, the overall ratio was set at 2.6, although for some simulations the 1990 rate of 2.31 was used instead. These figures, together with information from the comparison sample, provided the basic input into the Canadian evaluation's estimates of UI costs.

Overall, the Canadian evaluation found that STC cost the UI system approximately 35 percent more than an equivalent layoff alternative would have.[13,14] The researchers stated that the differences probably resulted primarily because 1) 29 percent of participating employees were laid off after their period of STC collection; 2) STC participation does not require the two-week waiting period, while regular UI does; 3) STC recipients were eligible for higher weekly benefit amounts than laid-off employees; and 4) UI is not collected by all eligible laid-off employees.

In summary, the Canadian study represents a hybrid in terms of the methodology used to measure the layoff conversion rate. The basic conversion rate used was primarily self-reported, but many adjustments were made to this rate with information from the study's relatively imperfectly matched control sample. It is interesting that the study yielded cost comparisons that are similar to those reported in the Best and MPR studies.

RESULTS FROM THE BPA/MPR STUDY

Our recently completed study of the STC program in the United States sheds additional light on the layoff conversion question, although it fails to provide a convincing numerical estimate of this parameter (Berkeley Planning Associates and Mathematica Policy

Research, Inc. 1997). As part of that project, we collected administrative data from a relatively large sample of firms that used STC during 1992 in five states (California, Florida, Kansas, New York, and Washington).[15] A comparison sample of equal size was also selected from these states, with firms being matched to the STC firms using three variables: firm size, three-digit industry, and UI tax rate. Because of the large number of non-STC firms in the states, we were able to match quite closely along these dimensions. Table 2 provides some quantitative measures of these matches. In general, the availability of a very large universe of firms that did not use STC made it possible to achieve a comparison sample that was virtually identical to the STC sample as measured by the available data. Still, as for all nonrandom evaluation methodologies, we were concerned that STC participants may have differed from nonparticipants along unmeasured dimensions—a concern that proved to be of crucial significance in the analysis of the data we collected.

Table 3 summarizes our data on regular UI and STC benefits charged to firms during the three years 1991, 1992, and 1993. To control for the large variation in firms' sizes in our sample, these data have been normalized by firms' total 1991 "full-time-equivalent payrolls" and then stated as a percent. Hence, a total normalized charges of 1.00 means that the firms' workers collected total chargeable benefits that amounted to 1 percent of the total 1991 payroll.[16]

Three general conclusions are immediately apparent from the table. First, although overall charges to the UI system varied among the three years, they were substantial in all three years, being somewhat larger in 1992 than in either 1991 or 1993. Second, regular UI was the predominate form of charges incurred. That finding was, of course, expected for firms in the comparison sample who, by definition, had no STC charges in 1992. The findings for firms in the STC sample are important, however. They show that, in 1992, between 62 and 78 percent of all UI system charges were for regular benefits; that is, despite their participation in the STC program, firms in 1992 appear to have used layoffs as their primary workforce reduction strategy. Thus, models that assume that firms follow an "either/or" approach to such strategies clearly are inappropriate.

Finally, and most important, the values in Table 3 suggests the possible difficulties in relying on a comparison methodology for address-

Table 2 Descriptive Statistics for Match Quality

Match characteristic	California[a]	Florida[b]	Kansas	New York[c]	Washington[d]
Correlation between number of employees in matched firms	0.98	0.84	0.70	0.9	0.79
Correlation between tax rates in matched firms	0.99	0.98	0.98	0.99	0.98
Number (percentage) of matches at the three-digit standard industrial classification (SIC) level	474 (93.5)	174 (82.1)	65 (63.7)	477 (94.5)	326 (87.2)
Number (percentage) of matches at the two-digit SIC level	27 (5.3)	20 (9.4)	28 (27.5)	27 (5.3)	44 (11.8)
Number of matched pairs of firms	507	212	102	505	374

SOURCE State administrative records.

[a] The 507 firms in California were selected from among 5,143 firms with STC plans in 1992 using sampling stratified by the number of employees and one-digit SIC code. There are 100 matches in California that are excluded from this table since the STC and comparison firms did not have complete information upon which a match could be made; that is, they were missing information on the number of employees or the tax rate.

[b] There are eight matches in Florida excluded from this table since the STC and comparison firms did not have complete information upon which a match could be made; that is, they were missing information on the number of employees or the tax rate.

[c] The 505 firms in New York were selected from among 737 firms with STC plans in 1992 using sampling stratified by the number of employees and one-digit SIC code. Only firms with at least five employees are eligible for participation in STC; comparison firms with fewer than five employees, therefore, were excluded from the pool of potential comparison employers.

[d] There are nine matches in Washington excluded from this table since the STC and comparison firms did not have complete information upon which a match could be made; that is, they were missing information on the number of employees or the tax rate.

Table 3 Average Compensated Unemployment Charges, STC and Non-STC Firms[a,b]

	California		Florida		Kansas		New York		Washington	
Characteristics	STC	Non-STC	STC	Non-STC	STC	Non-STC	STC	Non-STC	STC	Non-STC
1991										
Normalized UI charges	0.871	0.791	1.426	1.284	0.915	0.762	1.335	1.501	3.543	3.470
Normalized STC charges	0.277	0.007***	0.359	0.036***	0.343	0.027***	0.400	0.009***	0.368	0.039***
Normalized total charges	1.149	0.798**	1.785	1.320**	1.258	0.788***	1.735	1.510*	3.911	3.508
Percentage of total charges that are UI charges	79.424	98.531	84.060	98.773	77.536	98.387	71.393	99.433	90.275	98.519
1992										
Normalized UI charges	0.936	0.964	1.825	1.153***	1.681	1.206**	2.339	2.297	3.695	3.907
Normalized STC charges	0.561	0.000***	0.847	0.000***	0.759	0.000***	0.878	0.000***	1.022	0.000***
Normalized total charges	1.497	0.965***	2.672	1.153***	2.440	1.206***	3.217	2.297***	4.717	3.907***
Percentage of total charges that are UI charges	63.304	100.034[c]	62.078	100.000	69.030	100.000	65.146	100.000	78.456	100.000

1993										
Normalized UI charges	0.788	1.009***	0.935	1.088	1.366	1.137	1.783	1.810	3.633	4.298*
Normalized STC charges	0.331	0.131***	0.181	0.021***	0.258	0.040***	0.639	0.016***	0.893	0.084***
Normalized total charges	1.120	1.140	1.116	1.108	1.624	1.177*	2.421	1.826***	4.359	4.383
Percentage of total charges that are UI charges	74.999	92.423	79.442	99.564	85.169	98.894	69.354	98.438	81.856	97.039
Sample size	431	721	191	231	90	106	441	559	314	378

SOURCE: State administrative records.

[a] * = Significantly different from STC firms at the 10% level, two-tailed test.
** = Significantly different from STC firms at the 5% level, two-tailed test.
*** = Significantly different from STC firms at the 1% level, two-tailed test.

[b] Samples restricted to firms in business throughout 1991 and 1992. Because sample sizes vary slightly per charges measured, and because of rounding, the sum of normalized UI charges and normalized STC charges in a year may not equal normalized total charges in a year. All charges variables are normalized by an approximation of payroll at full employment in 1991. See text for further details.

[c] Firms occasionally have negative STC charges for a year. In these instances, the percentage of total charges that are UI charges may appear greater than 100%.

ing the layoff conversion question. In two of the states (Florida and Kansas), mean UI charges in 1992 were significantly larger in STC firms than in comparison firms. Because no current theory suggests that layoffs should be greater in otherwise similar firms that use STC, the most likely explanation of this result is that STC and comparison firms differ along dimensions that were not adequately controlled for in the matching process—for example, those firms which opt to use STC may have faced systematically worse economic prospects than did similar firms in the comparison group. Further support for this supposition is reflected in the figures for total charges in 1992, which were significantly larger for STC firms than for comparison firms in all of the states by amounts ranging from 0.53 percent of total payroll (California) to 1.52 percent (Florida).

We examined a number of statistical procedures for controlling for these unmeasured differences between the STC and comparison firms in models that seek to estimate the effect of STC on layoffs. For ease of comparison we present a variety of results for one simple specification:

$$(1)\quad Y_t = \alpha + \beta X_t + \gamma STC + \delta Y_{t-1} + u_t\,,$$

where Y is our measure of layoffs (normalized UI charges), X is a set of individual firm characteristics, STC is a dummy variable representing participation in the STC program, u is a random error term, and the lagged value of Y is included as a proxy for the general economic health of the firm in the previous year.[17,18] This specification was chosen both for its overall simplicity and because it was believed that estimates of γ, if they were unbiased, would permit a direct measure of what we have called the "layoff conversion rate."[19]

Table 4 reports results for four alternative approaches to the estimation of γ. The first specification used all observations in the sample for which we had complete data. As is immediately apparent, these estimates take on values that cannot reasonably be interpreted in the layoff conversion context. In three of the states, γ is estimated to be positive; in the other two, its value is very close to zero. One interpretation of such estimates is that they arise through selectivity bias—that is, unmeasurable characteristics that differ systematically between STC participants and firms in the comparison group cause estimates of γ to

Table 4 Estimated Effect of STC Participation on Normalized UI Charges in 1992 under Alternative Specifications[a]

Specification[b]	California	Florida	Kansas	New York	Washington
(1) Full sample					
STC dummy	–0.094	0.554***	0.487**	0.223	–0.075
Sample size	1,152	421	196	1,000	692
(2) Omit zero UI charges					
STC dummy	–0.301***	0.279	0.363	–0.101	–0.670**
Sample size	993	375	183	907	631
(3) Include zero STC charges					
STC dummy	–0.203	0.387**	0.546*	0.322*	–0.110
Sample size	1,152	416	194	1,000	692
(4) Matched pairs					
STC dummy	–0.073	0.574***	0.487*	0.233	–0.306
Sample size	502	336	132	814	444

SOURCE: U.S. Department of Labor Evaluation of Short-Time Compensation Programs. Administrative data.

[a] *Significantly different from zero at the 0.10 level, two-tailed test.
**Significantly different from zero at the 0.05 level, two-tailed test.
***Significantly different from zero at the 0.01 level, two-tailed test.

[b] All regressions contained an identical set of control variables for firm size and its square, UI tax rate, industry, and normalized UI and STC charges in 1991.

be biased in a positive direction. However, the first specification does control for some variables that might plausibly serve as proxies for the firms' economic health, such as their UI tax rates or their UI charges in the prior year. Hence, such selectivity must be based on current economic factors that are not adequately controlled for by these variables.

Because we did not have extensive information on other potential control variables, we employed a variety of sample restrictions to determine whether the suspected biases in specification (1) might be mitigated.[20] Our first approach focused on a possible asymmetry in the treatment definition. The STC treatment variable requires that some STC benefits were paid to a firm's workers in 1992, but no similar requirement was imposed for the comparison firms.[21] As an attempt to control for this asymmetry, we omitted from the comparison group all firms without UI charges in 1992. Results for this omission are shown in Table 4, as specification (2). Although this approach is admittedly ad hoc and runs the danger of incorporating biases of its own (an STC firm might not have made any layoffs in the absence of the program, and that decision would not be represented in such a sample), it did have a substantial effect on the estimates of γ. Both Kansas and Florida samples continued to exhibit (insignificantly) positive estimates, but estimates for the other three states were negative. In California and Washington, they were significantly different from zero.[22] These results clearly implied that selectivity is biasing parameter estimates in the full sample, but we had little confidence that our sample redefinition corrected for those biases in any systematic way.

Our final two specifications yielded similar ambiguous results. For specification (3), we redefined our STC variable to indicate only the filing of plans in 1992—it was not necessary for the firm to have experienced any benefit charges under the program. The rationale here was the converse of that employed for specification (2)—firms filing for STC were obviously aware of the program, but this specification required that neither STC firms nor comparison firms had actually incurred any charges in 1992. Again, however, this procedure yielded several significant positive estimates of γ. Finally, in specification (4), we returned to our original STC indicator variable but required that only pairs of STC firms and comparison firms enter the sample together. That procedure was intended to ensure that sample attrition because of insufficient data was not influencing our results.[23] Again, many of the estimates for γ were implausibly positive.

Our statistical examination of the layoff conversion issue therefore reached two primary conclusions. First, firms that use STC also make extensive use of layoffs. Even though participation in the STC program may offer significant advantages to these firms, they still appear to rely on compensated reductions in hours for much less than half of their total workforce adjustments. Further analysis of this outcome confirmed that STC use often accompanied widespread, "massive" layoffs (see Berkeley Planning Associates and Mathematica Policy

Research 1997, Chapter 6). Thus, it seems clear that modeling of the impact of STC-type programs on labor demand over the business cycle must take into account the full complexities of firms' actual workforce reduction strategies. Considerably more empirical work is required if we are to have reliable estimates of layoff conversion rates from STC use in various circumstances.

Our second primary conclusion is that adoption of a comparison methodology for the measurement of layoff conversions may be inadequate for obtaining unbiased estimates. Selectivity effects in program participation, at least for the low levels of participation experienced in the United States, pose major, perhaps insurmountable, problems in statistical inference. In our experimentation with alternative specifications, we obtained a wide variety of layoff conversion rate estimates, many of which seemed implausible on theoretical grounds. Indeed, our results suggest that the variation in earlier estimates of layoff conversion rates using matching methodologies (Kerachsky et al. 1986) may also be in part explained by selectivity biases. Further progress on estimation of this parameter may well require the development of alternative statistical approaches.

CONCLUSIONS

In theory, the availability of STC benefits in addition to benefits provided through regular unemployment compensation should affect how firms make cyclical workforce adjustments. Understanding the quantitative magnitude of these effects is an important component in any evaluation of a program's overall desirability. Given the centrality of this issue, we find it surprising that relatively little attention has been paid to the specification of a clear model for estimating these effects. In our view, estimation of this type of model should be in the forefront of economic research on STC programs.

Our review of the empirical literature on STC suggests that how program availability affects firms' workforce dynamics is far from clear. Consistent with other evaluations of their type, studies based on self-reporting have produced widely varying estimates, some of which are implausible. Many other studies merely assume the size of an effect that should, in principle, be estimated. Matching methodologies seemed to offer the best promise of obtaining estimates of the effect of STC from microdata. Previous experiences with that methodology in other contexts, however, suggest that its greatest drawback is the lack of assurance that STC users and nonusers face similar economic prospects despite major efforts to assure that the firms are closely matched on measurable variables. This possibility was confirmed by our experiences in trying to model the impact of firm use of STC in the United States during the 1992 recession, for which we obtained a wide variety of layoff conversion rate estimates. Hence, the suitability of other research designs (such as innovative uses of aggregate data, random-assignment experiments, or carefully designed case studies) needs to be considered before attempting any overall assessment of the general desirability of STC-type programs.

Notes

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1. Utilization rates for STC have been low under both programs, however. UI programs in the United States also have provisions for the payment of partial benefits, but these benefits are generally unavailable to workers suffering relatively modest workweek reductions.
2. Studies of the increase in short-time work during recessions provide additional support for the possibility that total compensation might increase when STC becomes available.
3. In principle, one might want to measure changes in total hours and employment in response to STC use, not simply unemployed hours that are compensated. But the data requirements for a more complete measurement at the level of the firm are quite onerous, and no researcher has attempted such an evaluation. Rather, the existing research has focused on more readily measured compensated hours, usu-

ally by drawing data from administrative sources. Although use of the compensated data sheds light on a number of important policy questions (such as the effect of STC adoption on overall expenditures under the UI system), the extent to which these data accurately reflect changes in total hours and employment is not known.

4. For summaries of many of the other issues that have been addressed in the STC research, see Best (1988) or Cook et al. (1995).
5. A firm may choose not to use its approved plan. State agencies that approve plans do not typically monitor whether plans are used and what happens if they are not used.
6. Other assumptions pertain to the UI take-up rate of STC participants, the hazard rate to reemployment for laid-off employees, the wages and benefit levels of STC participants, and the costs of processing STC claims. The estimate assumes firms operate at the average values of workweek reductions, STC durations, and post-STC layoffs.
7. In some cases, effects of STC on aggregate fluctuations make no use of the STC data and therefore need no assumption about layoff conversions. For example, Van Audenrode (1994) finds that total hours exhibit much greater flexibility in European countries with generous STC systems (Belgium, Italy, and Sweden) than in countries with only modest levels of compensation (France and Germany). Although this finding does not provide a direct estimate of the extent to which STC use deters layoffs, it suggests that the trade-off may be significant in certain situations.
8. Of course, the reason firms choose not to use STC is of critical importance. Firms that were not aware of STC might be more suitable matches than firms that knew about it and chose not to use it.
9. Economic or operational criteria, such as legislative restrictions on firm characteristics that limit eligibility for STC, may suggest the need to exclude certain firms. In addition, there may be no credible matches for a particular firm (see, for example, Kerachsky et al. 1986; and Schiff 1986).
10. Even if additional data were available for matching, the matching procedure is computationally burdensome and extremely slow.
11. The authors made no attempt to reweight the sample to adjust for the higher probability of sampling large firms.
12. The analysis on pp. 148–155 indicates that STC claimants collected benefits for fewer weeks than UI claimants, but it appears that adjustments were not made to account for the increased number of claimants under an STC program. Furthermore, compensation under STC was only for the workweek reduction and not for the full weekly benefit amount.
13. Canada's UI system is not experience rated, so higher compensated unemployment charges are not charged to a firm. Because the U.S. system is experience rated, albeit imperfectly and with a lag, firms with higher charges typically bear responsibility for them. Hence, computations of the U.S. program's "cost" to the UI system must, of necessity, be more complex.

14. For firms, the evaluation's reported favorable benefit–cost estimates for STC compared to layoffs resulted largely from the significant savings derived from reduced training and hiring costs. For society, the favorable estimates resulted largely from the much lower stress-related costs of STC compared to layoffs. These values were calculated on a per-layoff equivalent basis using the self-reported figures described earlier.
15. Our definition of having "used" STC was that some benefits were paid to the firm's workers during 1992 under this program. Implementing this definition posed some difficulties because the only information available was on which firms had filed STC plans at the time of sample selection. Hence, we selected our STC sample on the basis of having filed a plan in 1992, although (as discussed later) we primarily used a definition stressing actual use in most of our analysis.
16. 1991 full-time-equivalent payrolls were estimated by adding total wages paid during 1991 to an estimate of wages that would have been paid to the firms' workers who collected UI or STC during 1991 if these workers had instead been fully employed during these periods.
17. The vector X includes firm size and its square, dummy variables for one-digit SIC industry, and a measure of the 1991 UI tax rate.
18. Prior year normalized STC charges were also included in all regressions as a further control on the firm's health in 1991.
19. Specifically, the average layoff conversion rate can be estimated as γ/k, where k represents mean normalized STC charges during 1992. In this specification, the layoff conversion rate would be in terms of dollars—STC dollars substituting for UI dollars. We obtained substantially similar results from estimates of Y (and γ) in terms of hours—perhaps a more natural, if less accurately measured, conversion concept.
20. We also experimented with a number of statistical methods for controlling for sample selectivity, but these were largely unsuccessful because of our inability to develop clear ways of identifying the selectivity equation given the limited set of control variables we had and the fact that firms in the sample had been matched on these variables.
21. This asymmetry was also pointed out in connection with the earlier MPR study (Morand 1990).
22. The point estimates for California and Washington imply layoff conversion rates of –0.55 and –0.66, respectively.
23. To ensure the integrity of the matching, pairs in which a comparison firm experienced STC charges (because of a plan filed prior to 1992) were also omitted from the sample in some specifications. We also used this sample to implement a "paired" regression analysis, but this did not substantially change the results.

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10

Short-Time Work in the United States

Implications for Evaluation of Short-Time Compensation Schemes

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Abraham and Houseman's (1993, 1994) calls for a systematic reorganization of U.S. job security policies have again drawn attention to reduced-hours employment—financed through short-time compensation (STC)—as an alternative to layoffs. Under STC,[1] workers receive partial unemployment insurance (UI) benefits as compensation for reduced working hours. While STC programs are widely used in countries such as Belgium, Canada, France, and Germany, in the United States they have been available in only 18 states, and the usage rates are very low (Vroman 1990, 1992).[2]

The literature typically assumes that workers receiving STC are not drawn from the pool of those who would have had their hours reduced anyway, even without an STC program. Such reduced-hours employment with no offsetting compensation is called short-time work (STW). There has been virtually no discussion of the role of STW in evaluating STC programs in the existing literature. This chapter analyzes the incidence of STW in the United States for 1968–1993 and the potential implications for evaluating the impact of STC programs.

Conventional wisdom holds that STW is not widely used in the United States. Yet the answer depends on the measure of STW used. In terms of incidence per worker, STW use is fairly widespread: the STW rate ranges from about 50 to 70 percent of the layoff rate.[3] However, in terms of total hours adjustment, the STW rate is dwarfed by the layoff rate.

The main implication for STC program evaluation is that widespread expansion of STC programs is most likely to benefit workers already on or subject to STW, not necessarily workers on or subject to layoffs. There is not necessarily anything wrong with this conclusion. Just as the UI system was established to provide consumption insurance to workers subject to employment shocks, STC programs could be expanded to insure STW workers against partial employment shocks. However, this aspect of STC has not been a focus of previous authors, who instead have concentrated on the spillover effects of STC on layoffs.

The chapter is organized as follows: the next section discusses the existing literature on STC and related evidence on employment adjustment using layoffs versus hours; the following section presents the empirical evidence that STW is already widely used in the United States with particular attention paid to the prevalence of STW in STC and non-STC states; and the last section concludes the chapter.

SHORT-TIME COMPENSATION AND SHORT-TIME WORK

Unemployment insurance was conceived as a program to insure workers' consumption against unexpected employment shocks. It was presumed that the beneficiaries would be people who would suffer an unemployment spell regardless of whether there was a UI payment system. However, an extensive literature has highlighted an unintended side effect of the UI system in the absence of perfect experience rating: workers get laid off who would otherwise have remained employed because UI payments offer an implicit subsidy to layoffs (Hamermesh 1993, pp. 307–315). Many authors have attempted to measure the number of "excess" layoffs created by imperfect experience rating. Approximately 20–40 percent of temporary layoffs are in this category (Topel 1983; Card and Levine 1994; Anderson and Meyer 1994). This figure is substantial and has implications for potential overuse of STC, as discussed below. However, the pertinent observation for this section is that a majority of UI claimants would have been laid off even in the absence of a UI program. This suggests that the primary beneficiaries of STC programs are likely to be people who would be put on STW even in the absence of an STC program.

There are three pools of workers from which STC-compensated employees can be drawn. Each group is denoted by the state in which they would have been in the absence of an STC program: workers who are laid off in response to a demand shock, workers whose hours are not adjusted at all, and workers whose hours are cut back. The prevalent view in the literature is that only the first and second groups of workers are tapped for inclusion in STC programs:

> [S]hort time compensation (STC) represents an alternative to layoffs as a way for firms to reduce labor inputs in periods of slack demand. Currently the standard procedure for reducing work hours is to lay off the least senior employees. This action concentrates the reduction in hours narrowly among a small number while leaving other workers unaffected. An alternative procedure for reducing labor input is to retain all employees by reducing weekly hours for a much larger fraction of the firm's work force. (Vroman 1990, p. 71)

Little mention is made of the existence of STW in the absence of STC, particularly by those advocating STC programs. One exception is Hamermesh (1978, pp. 249–250), who noted that

> While the subsidy [STC] will to some extent encourage the expansion of the activity that is subsidized [STW], it will also reward those economic agents—in this case firms and workers—that would engage in the subsidized activity even in the absence of the subsidy. Because of this windfall much of the payment for short-time work under any STC scheme cannot result in increased employment, but is instead a transfer from those whose taxes exceed their receipts from STC to those for whom the opposite is true.

The vast majority of authors since Hamermesh have simply ignored this issue. One of the lone exceptions is Best (1981, p. 96), who dismisses Hamermesh's critique, stating that

> [T]he incidence of such workweek reduction [STW] appears to be low in the United States and commonly smaller than the 10 percent threshold reduction of worktime required before employees are eligible to receive benefits.

In support of this conclusion, Best cites only one source, Henle (1978, p. 267), who stated that "the evidence seems to indicate that the prevalence of such work sharing arrangements is quite limited." This conclusion was based on union contract data showing that about 20 percent of contracts provided for hours reductions in the face of slack work, and that such provisions were generally not utilized. However, this was the full extent of statistics provided to support these conclusions.[4] Moreover, because unionized workers account for only a minority of the workforce, Best's dismissal of the importance of STW is clearly premature without additional evidence for nonunionized workers.

Bednarzik (1980), the only other author to analyze the incidence of STW, tracked the aggregate STW rate for 1956–1979. However, the only comparison made with other aspects of the labor market was the aggregate unemployment rate. This created the impression that STW is relatively underutilized because the unemployment rate is many times larger than the STW rate. The proper comparison for an STC evaluation is STW versus layoffs, because both represent employer-initiated changes in hours in response to demand shocks. Moreover, layoffs typically account for only about 15 percent of unemployment in any given year (Economic Report of the President 1996). As shown below, the incidence of STW in fact is often comparable to the incidence of layoffs.

EMPLOYMENT VERSUS HOURS ADJUSTMENT

It has long been recognized that fixed costs of hiring and firing workers inhibit a firm from using employment adjustment as the only way to adjust total labor input (Oi 1962). Consequently, in the short run firms adjust hours per worker as a substitute for adjusting the number of workers (Rosen 1968; Fair 1969; see Caballero et al. 1995 for a recent example using data from individual manufacturing plants).

Previous research has directly compared employment and hours adjustment for the United States versus other countries that have much more liberal STC provisions (Abraham and Houseman 1993; Van Audenrode 1994). Both Abraham and Houseman and Van Audenrode found that the adjustment of total hours in the United States is done more through employment than through average hours per worker. Both studies concluded that the relative lack of a generous STC subsidy plays a role in this: U.S. firms use employment adjustment relatively more than hours adjustment presumably because the former are more heavily subsidized. However, Van Audenrode suggested that a reduced reliance on layoffs would occur only if the proportionate subsidy to STC exceeded that for layoffs. Abraham and Houseman advocated an expansion of STC at the same time that experience rating for layoff UI benefits is tightened. Thus, both studies did not presume that simple changes to STC alone would necessarily reduce firms' reliance on layoffs.

These authors' hesitance to advocate expansion of STC as the only way to shift labor adjustment away from layoffs is well-founded. In particular there are both institutional and mechanical differences between labor markets across different countries. Though STW may be used relatively more in countries with more liberal STC, such a correlation is not proof that changes in STC provisions would produce a similar reliance on STW in the United States. In particular, tighter experience rating of UI alone might eliminate the excess reliance on layoffs without the need for a generous STC subsidy. This is precisely the point made by Burdett and Wright (1989), who show that an STC subsidy leads to an inefficient number of hours per worker.

The reliability of cross-country comparisons such as those above are also limited by the nature of shocks that hit particular industries. Aggregate net employment changes mask much larger offsetting flows through gross job creation and destruction. In particular, there are large differences in job reallocation rates between countries (Davis, Haltiwanger, and Schuh 1996, p. 21). It is naive to presume that such cross-country differences can be fully explained by parameters of UI and STC alone. They undoubtedly arise due to differences in a host of factors such as country size, population and industrial concentration, internal migration patterns, barriers to entry for new businesses, merger and takeover rules, bankruptcy laws, union organizing laws, the demographic makeup of the labor force, societal differences in between-job mobility, welfare system influences on work behavior, overtime pay rules, etc.[5]

Limiting the analysis to nominally comparable, narrowly defined industries (as do Abraham and Houseman 1993) does not negate the role of these other factors that affect the ability and preferences of individual firms to adjust labor input. Abraham and Houseman (1994) partially address this issue by analyzing the effect of weakening employment security laws in Germany, France, and Belgium. They find that such changes—which presumably decreased the costs of layoffs—did not measurably increase reliance on employment (over hours) adjustment in those countries. While informative, such evidence is not proof that expanding STC coverage in the United States would increase reliance on hours adjustment. If anything, their results suggest that such an expansion could easily have no measurable impact on the use of layoffs.

Thus, there is a clear need to analyze the use of STW in the United States as a way of predicting the impact of STC programs. Such an analysis is better than cross-country comparisons of employment versus hours adjustment because a vast majority of between-country differences in other factors are held constant for a within-country analysis. Moreover, the data used here allow the identification of hours reductions below usual hours worked.[6] They also allow the identification of employment adjustment through layoffs. Both of these are more accurate measures of the relevant margins on which firms actively decrease labor input than measures such as the relative usage of total hours adjustment versus total employment adjustment (which include both increases and decreases in labor input). In particular, the latter measures include labor turnover that occurs through hiring and voluntary separations, which are important components of labor adjustment but are not of primary importance for predicting firms' responses to changes in layoff versus STW subsidies.

THE USE OF STW IN THE UNITED STATES

The data used for this study are drawn from two different sets of Current Population Survey (CPS) data: the 1968–1993 March Annual Demographic Files, and the 1979–1993 Outgoing Rotation Group Files (for all 12 months in the year). The sample was limited to wage and salary workers age 16 and older. A worker is defined to be on STW if 1) the total number of hours worked during the survey week (at all jobs) are less than 35, 2) usual hours are greater than 35, and 3) the reason given for working less than 35 hours during the survey week is slack demand, material shortage, or plant/machine repair. Bednarzik (1980) included only those who indicate slack work in his measure of STW. I include the other two because they also represent employer-initiated hours reductions, which could be induced by demand shocks. Regardless, these two categories consistently account for less than 10 percent of STW, so excluding them would not substantively alter any of the conclusions.

The analysis covers data through 1993 because data for 1994 and after are not directly comparable. The U.S. Bureau of Labor Statistics (1994) introduced a major redesign of and improvement to the CPS in 1994, making comparisons with earlier years problematic. A full reconciliation of the STW rates for 1993 and earlier compared with 1994 and later is beyond the scope of this study. However, one important difference should be noted. During 1993, the U.S. Bureau of Labor Statistics administered both the old and new versions of the CPS. This allowed for a single year comparison of differences in the measured levels of layoffs and STW.

People on STW are a subset of those who usually work full time yet are part time during the survey week. According to the new, more accurate survey, the actual level of this larger category of "temporary" part-time workers—of which STW is a subset—is 25 percent lower than recorded in the old survey. In contrast, the actual level of layoff unemployment is 10 percent higher (U.S. Bureau of Labor Statistics 1994). Unfortunately, there is no way to determine definitively whether these biases in the old measures (the ones used in the present study) were consistent over time. However, there is no particular reason to believe they were not consistent. Regardless, the reader should keep in mind the biases while reviewing the empirical evidence below.

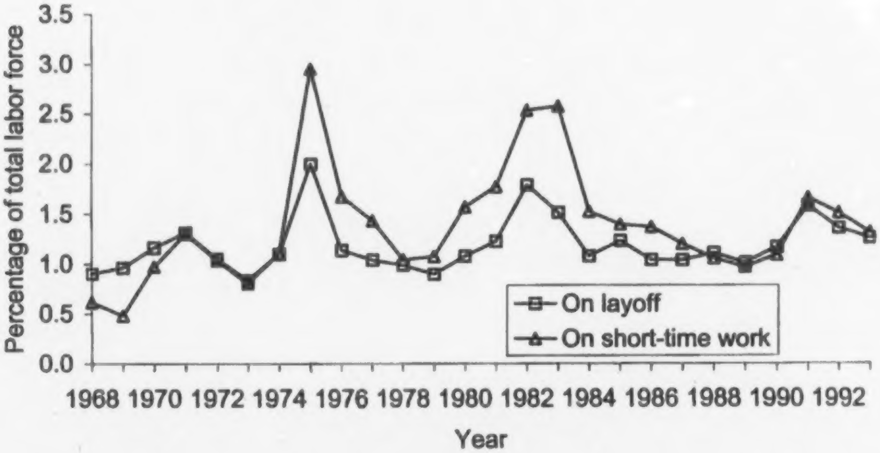
CYCLICAL PATTERNS

Figure 1 graphs the rates of STW and layoffs for 1968–1993 using the March data. Throughout the chapter, the STW and layoff rates are calculated using the same base for the labor force: all those employed or on layoff. People unemployed for reasons other than layoff are excluded from the analysis.[7]

The incidence of STW in Figure 1 is comparable to layoffs; the layoff rate is appreciably higher only during the 1970s and early 1980s recession years. Throughout the most recent recession, the STW and layoff rates were virtually identical. Adjusting for the biases mentioned above would raise the layoff rate by 10 percent and lower the STW rate by 25 percent. Accounting for this, the true STW rate is approximately 50 to 70 percent of the layoff rate. The use of STW is concentrated in industries such as construction that heavily use layoffs, as shown in Table 1.

While the incidence of STW is comparable to layoffs, total hours adjustment is comparable only to short-term layoffs. This can be seen in Figure 2, which graphs the incidence of STW and layoffs in the top panel (as a fraction of employment), and the percentage of total hours

Figure 1 The Incidence of Layoffs and Short-Time Work for Entire Labor Force



SOURCE: Calculations based on March Current Population Surveys, 1968–1993.

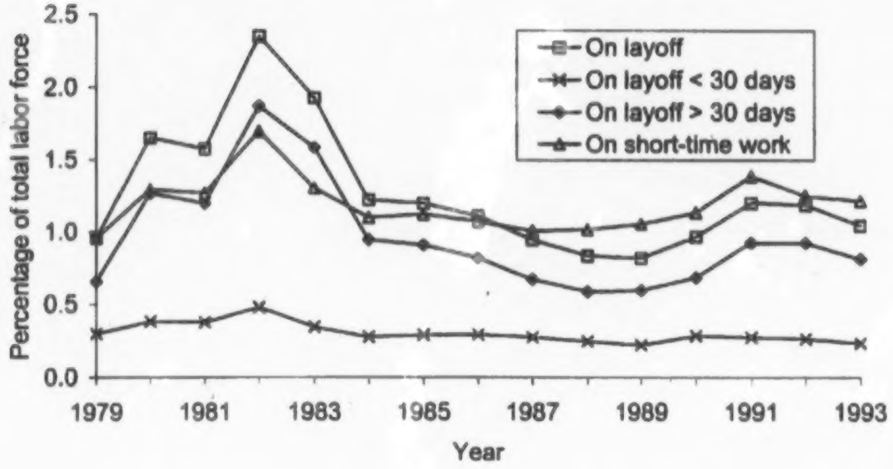
Table 1 Short-Time Work Rates as a Share of Employment and of Total Hours[a] (%)

			Selected industries						Selected occupations			
	Total labor force		Construction		Durable manufacturing		Public administration		Clerical/ administration support		Skilled laborers	
Year	Emp.	Hrs.	Emp.	Hrs.	Emp.	Hrs.	Emp.	Hrs.	Emp.	Hrs.	Emp.	Hrs.
1979	0.96	0.33	2.65	1.08	0.84	0.31	0.15	0.05	0.38	0.12	1.78	0.64
1980	1.29	0.43	3.37	1.43	1.53	0.47	0.28	0.09	0.45	0.13	2.50	0.86
1981	1.27	0.41	3.48	1.50	1.35	0.46	0.15	0.07	0.39	0.13	2.44	0.81
1982	1.69	0.52	4.37	1.68	2.58	0.73	0.16	0.06	0.74	0.21	3.52	1.15
1983	1.30	0.40	3.37	1.27	1.50	0.42	0.11	0.04	0.52	0.15	2.71	0.88
1984	1.10	0.38	2.84	1.18	1.01	0.32	0.15	0.07	0.42	0.15	2.11	0.76
1985	1.12	0.35	2.69	1.09	1.21	0.36	0.13	0.03	0.42	0.12	2.28	0.75
1986	1.07	0.33	2.75	1.11	0.94	0.29	0.16	0.05	0.44	0.12	1.96	0.69
1987	1.01	0.33	3.17	1.25	0.79	0.25	0.15	0.05	0.41	0.11	1.79	0.66
1988	1.02	0.32	2.94	1.03	0.91	0.29	0.19	0.05	0.40	0.11	1.93	0.67
1989	1.05	0.33	3.21	1.09	0.92	0.32	0.22	0.08	0.54	0.14	1.87	0.66
1990	1.13	0.35	3.47	1.33	0.95	0.31	0.15	0.05	0.46	0.14	2.02	0.68
1991	1.38	0.43	4.81	1.86	1.48	0.43	0.12	0.06	0.65	0.19	2.61	0.91
1992	1.25	0.38	4.24	1.63	1.05	0.32	0.21	0.07	0.51	0.15	2.29	0.77
1993	1.21	0.36	3.56	1.26	0.83	0.28	0.14	0.06	0.53	0.16	1.97	0.68

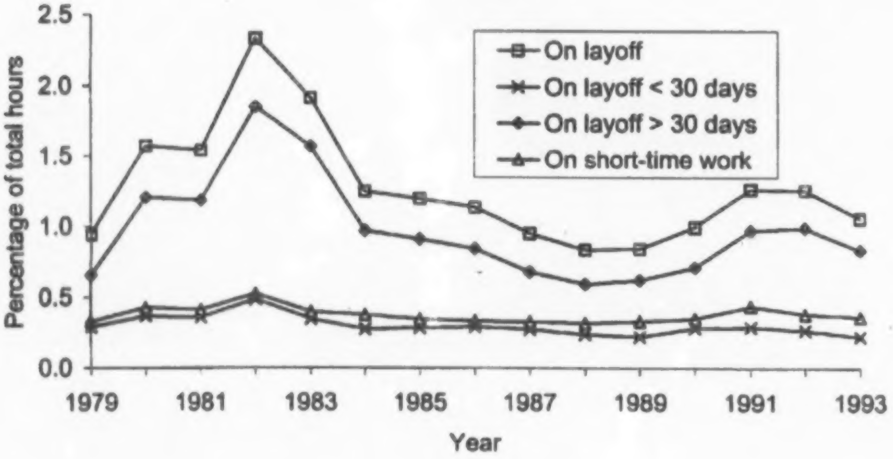
[a] The STW rates were calculated over all workers plus unemployed in each category using CPS Outgoing Rotation Group data for all months in the year.

Figure 2 Layoffs and Short-Time Work: Incidence versus Hours

Incidence



Hours Adjustment



SOURCE: Calculations based on Current Population Surveys, Outgoing Rotation Groups (all months), 1979–1993.

adjusted through both channels in the bottom panel. For the bottom panel, usual hours for persons on layoff had to be imputed because the CPS does not record that measure for people not with a job.[8] In both panels, total layoffs are broken down into two separate groups: those of duration less than 30 days and those of greater duration. Note that this refers to ongoing duration as of the survey date, so a significant portion of the short-term layoffs ex post will be longer than 30 days. But such a division is useful because the short-duration category undoubtedly includes a disproportionate number of layoffs that ex post will be less than 30 days.

The data in Figure 2 and throughout the rest of the chapter, use the Outgoing Rotation Group files so that the numbers are indicative of employment behavior for the entire year, not just March. This limits the time series to 1979–1993. However, Figure 1 shows that the degree of cyclical correlation between STW and layoffs barely differs for 1968–1978 versus 1979–1993. So the analysis for the most recent years should provide results comparable to the earlier period. Moreover, the overall pattern in the incidence of STW and layoffs in Figure 1 and the top panel of Figure 2 are virtually identical, showing that the year-to-year movements in the two rates in Figure 1 are not contaminated by cyclical factors that are unique to March.

The graphs in Figure 2 show that 1) the incidence of STW is comparable to all layoffs, particularly since 1987, yet 2) total hours adjustment through STW is only a fraction of total hours adjustment through layoffs. This is not surprising when one considers the likely source of demand shocks inducing the different types of adjustment. Firms that put workers on STW or on layoff for a short period of time probably have been hit by what are perceived to be temporary demand shocks. In contrast, firms with workers who have been on layoff for more than a month probably have been hit by what are perceived to be more permanent demand shocks.

This suggests that STW is more likely a substitute for short-duration layoffs than for long-duration layoffs. If a firm needs to downsize permanently, providing a short-term subsidy to STW through STC should not induce the firm to retain more workers in the long run. An STC subsidy might temporarily postpone such layoffs, if at all, but Figure 2 suggests that such a postponement may be quite short.

STW BY INDUSTRY AND BY OCCUPATION

Tables 1–3 report the pattern of STW usage by industry and by occupation. Tables 1 and 2 report incidence and hours measures for select industries and occupations by year and by month, respectively. Regardless of which measure is used, the more highly cyclical industries, such as construction and manufacturing, and the more highly cyclical occupations, such as skilled laborers, have the highest rates of STW. However, as seen in Table 2, there is a distinct seasonal pattern in STW for construction, with the highest rates in the winter and early spring.[9] The seasonal pattern for durable manufacturing is much less pronounced. This provides further evidence that usage of STW mirrors that of layoffs.

Consolidating all the data for 1979–1993, Table 3 examines which industries and occupations use STW the most. As foreshadowed by the patterns in Tables 1 and 2, the highest rates are for those that are the most cyclical and/or seasonal: the manufacturing, construction, mining and agriculture industries, and the skilled laborer, semiskilled laborer, and farming occupations.

Best's comment about the rate of reduction of hours under STC (". . . such workweek reduction appears to be . . . commonly smaller than the 10 percent threshold reduction of worktime required before employees are eligible to receive benefits") can be assessed by analyzing the STW incidence versus hours rates in Table 3. An estimate of the average hours reduction under STW is available by taking the ratio of the hours adjustment figure in column 5 over the employment adjustment figure in column 2.[10] Doing so yields an average reduction in hours of about 30 percent for each industry and occupation. As shown in Table 4, this figure falls well within the range necessary to trigger eligibility for STC for all states with such a program.[11] So Best's statement appears to be inaccurate by this measure, at least for current STC programs.

Table 2 Short-Time Work Rates as a Percentage of Employment and as a Percentage of Total Hours, Disaggregated by Month[a]

			Selected Industries						Selected Occupations			
	Total labor force		Construction		Durable manufacturing		Public administration		Clerical/ administrative support		Skilled laborers	
Month	Emp.	Hrs.	Emp.	Hrs.	Emp.	Hrs.	Emp	Hrs.	Emp	Hrs.	Emp	Hrs
Jan	1.39	0.46	3.74	1.60	1.29	0.42	0.12	0.05	0.53	0.17	2.51	0.90
Feb	1.32	0.42	3.62	1.56	1.42	0.43	0.17	0.05	0.50	0.14	2.53	0.89
Mar	1.22	0.39	3.52	1.46	1.37	0.42	0.12	0.08	0.51	0.14	2.51	0.85
Apr	1.13	0.37	3.74	1.44	1.20	0.42	0.20	0.06	0.47	0.16	2.28	0.80
May	1.16	0.37	3.26	1.31	1.36	0.41	0.11	0.05	0.46	0.14	2.31	0.79
June	1.18	0.36	3.24	1.22	1.11	0.31	0.26	0.07	0.53	0.15	2.24	0.73
July	1.20	0.37	3.12	1.21	1.07	0.32	0.17	0.06	0.56	0.16	2.10	0.70
Aug	1.11	0.34	2.66	1.00	1.12	0.34	0.16	0.05	0.47	0.14	2.08	0.66
Sept	1.04	0.32	2.84	1.04	0.98	0.33	0.28	0.08	0.42	0.12	1.87	0.66
Oct	1.10	0.34	3.25	1.20	1.13	0.35	0.21	0.07	0.40	0.11	2.05	0.69
Nov	1.18	0.36	3.52	1.23	1.27	0.39	0.07	0.02	0.44	0.14	2.30	0.78
Dec	1.28	0.42	3.96	1.54	1.19	0.36	0.13	0.06	0.50	0.14	2.39	0.85

[a] The STW rates were calculated over all workers plus unemployed in each category using CPS Outgoing Rotation Group data for 1979–1993.

Table 3 Short-Time Work and Layoff Rates by Industry and Occupation, 1979–1993[a]

	As a % of employment			As a % of total hours		
	Layoffs ≥30 days	STW	Layoffs <30 days	Layoffs ≥30 days	STW	Layoffs <30 days
Industry						
Agriculture	1.95	3.09	0.52	1.77	1.22	0.44
Mining	2.69	1.42	0.62	2.63	0.62	0.65
Construction	3.45	3.36	1.10	3.54	1.31	1.11
Durable manufacturing	2.15	1.21	0.57	2.12	0.37	0.57
Nondurable manufacturing	1.34	2.37	0.60	1.29	0.71	0.58
Transportation	0.80	0.88	0.22	0.80	0.36	0.22
Wholesale trade	0.72	0.74	0.18	0.68	0.26	0.16
Retail trade	0.53	1.18	0.18	0.51	0.24	0.16
Services	0.36	0.75	0.13	0.35	0.21	0.12
Public administration	0.26	0.17	0.05	0.25	0.06	0.06
Occupation						
Clerical/ administration	0.46	0.48	0.12	0.46	0.14	0.12
Skilled laborers	2.22	2.26	0.72	2.19	0.77	0.72
Educators	0.16	0.32	0.05	0.17	0.12	0.05
Farming, forestry	1.84	2.75	0.52	1.85	1.15	0.48
Medical/health	0.21	0.68	0.09	0.19	0.18	0.09
Management-related	0.27	0.29	0.04	0.27	0.11	0.04
Semiskilled laborer	2.53	2.65	0.76	2.66	0.89	0.79
Profess. specialty, not elsewhere classified	0.33	0.37	0.09	0.34	0.13	0.08
Personal service	0.69	1.71	0.27	0.69	0.34	0.25
Private household service	0.12	1.27	0.24	0.15	0.19	0.26
Protective service	0.38	0.40	0.11	0.37	0.12	0.09
Sales-related	0.43	0.89	0.14	0.43	0.19	0.12

[a] The STW and layoff rates were calculated over all workers plus unemployed in each category using CPS Outgoing Rotation Group data for all months during 1979–1993.

Table 4 Summary of Short-Term Compensation Programs[a]

Participating states	Duration of plan before new approval required	Limits on number of weeks	Required reduction of work (%)
Arizona	1 year	26	10 to 40
Arkansas	1 year	26	10 to 40
California	6 month	[b]	at least 10
Connecticut	6 month	26[c]	20 to 40
Florida	1 year	26	10 to 40
Iowa	2 year	26	20 to 50
Kansas	1 year	26	20 to 40
Louisiana	1 year	26	20 to 40
Maryland	6 month	26	10 to 50[d]
Massachusetts	6 month	26	10 to 60
Minnesota	1 year	52	20 to 40
Missouri	1 year	26	20 to 40
New York	—	20	20 to 60
Oregon	1 year	26	20 to 40
Rhode Island	1 year	26	10 to 50
Texas	1 year	52	10 to 40
Vermont	6 month[e]	26	20 to 50
Washington	1 year[e]	26	10 to 50

SOURCE: National Foundation for Unemployment Compensation & Workers' Compennsation (1996).

[a] As of January 1996.

[b] No limit on number of weeks, but total paid can not exceed 26 × weekly benefit amount.

[c] 26-week extension possible.

[d] 50 percent maximum may be waived by Secretary.

[e] Or date of plan, if earlier.

STC VERSUS NON-STC STATES AND PARTIAL UI

Are STC programs the reason for the existence of STW in the United States? In order to provide a crude answer to this question, Table 5 reports STW rates for states that have never had an STC program ("non-STC states") and states that have ever had an STC program ("STC states," including Illinois). The answer is no, because the STW rate has been greater for non-STC states in all years except for 1992–1993. It is true that many STC programs have only been introduced recently (Table 4), so the higher rates of STW in the most recent years for the STC states may be due to the recently adopted STC programs.[12] However, Figure 3 shows that that would be a premature conclusion: both STW and layoffs have been relatively higher in STC states in recent years, suggesting that both are correlated with other factors, such as different industrial compositions in the two groups of states. So it is doubtful that usage of STC explains much of the difference in STW between STC and non-STC states. This is consistent with the commonly held view that STC programs have been vastly underutilized in the states that have them.

One component of the UI system in each state that may explain at least part of the cross-state variation in STW is partial UI benefits. Partial UI benefits are available in all states when earnings fall below a particular threshold. However, the threshold is so low that the workweek has to be reduced by at least 60 percent in most states. This fact alone indicates that the provisions for partial UI probably are not a major factor in determining STW because the average hours reduction is only half that needed to qualify for partial UI. Moreover, in most states the partial UI benefit is taxed at a 100 percent rate for any earnings above a very small amount (the "disregard" amount).

Despite the fact that partial UI probably is not generous enough to explain patterns of STW, a crude test is provided in Table 6, which breaks the non-STC states into three groups: least generous, more generous, and most generous partial UI benefits. States in the first group tax partial UI benefits at a 100 percent tax rate for any earnings above either 10 percent of wages or $40 per week. States in the second group also tax benefits at a 100 percent tax rate for earnings over the disregard, but the disregard is higher than for the first group. The third

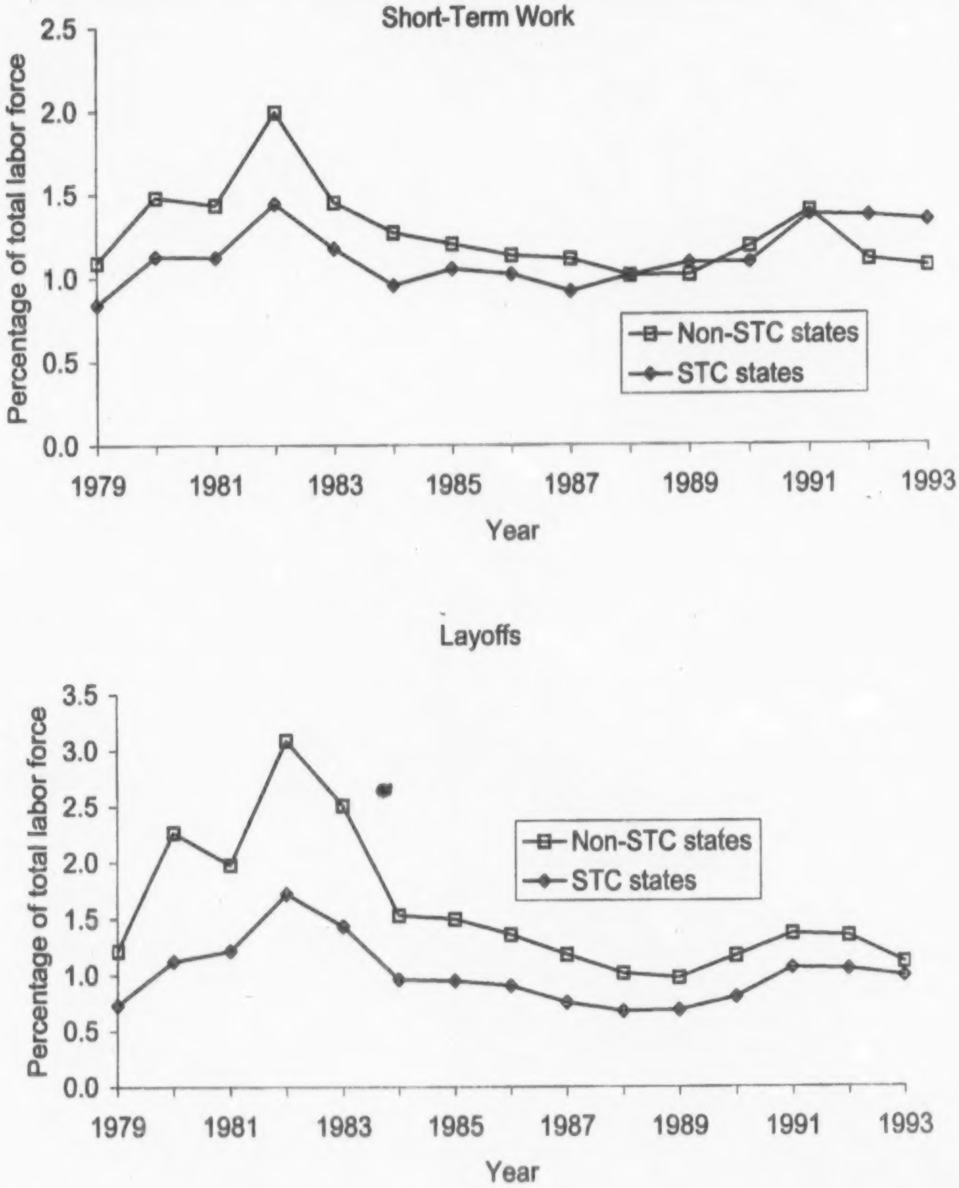
Table 5 Short-Time Work Rates as a Share of Employment for States with and without STC Programs[a] (%)

				STC states[b]			
Year	All states	All non-STC states	All	California	New York	Kansas	Missouri
1979	0.96	1.09	0.84	0.97	0.57	0.54	1.12
1980	1.29	1.48	1.13	1.34	0.73	0.86	1.24
1981	1.27	1.44	1.12	1.39	0.61	0.64	1.47
1982	1.69	1.99	1.44	1.78	0.78	0.99	1.55
1983	1.30	1.45	1.17	1.38	0.87	0.94	1.69
1984	1.10	1.27	0.96	1.03	0.67	0.70	1.16
1985	1.12	1.20	1.05	1.17	0.53	0.74	0.72
1986	1.07	1.13	1.02	1.13	0.66	1.19	1.07
1987	1.01	1.11	0.92	1.03	0.50	0.55	1.07
1988	1.02	1.01	1.01	1.21	0.55	1.10	1.58
1989	1.05	1.01	1.09	1.33	0.63	1.04	0.98
1990	1.13	1.18	1.09	1.57	0.69	0.56	1.13
1991	1.3	1.39	1.37	2.01	1.00	0.78	1.27
1992	1.25	1.10	1.36	2.17	0.96	0.45	1.52
1993	1.21	1.06	1.33	2.04	1.10	0.87	0.76

[a] The STW rates were calculated over all workers plus unemployed in each category using CPS Outgoing Rotation Group data for all months in the year.

[b] The STC states category includes all states that have ever had an STC program, even if the program was not in existence for one or more years during 1979–1993: Arizona, Arkansas, California, Connecticut, Florida, Illinois, Iowa, Kansas, Louisiana, Massachusetts, Maryland, Minnesota, Missouri, New York, Oregon, Rhode Island, Texas, Vermont, and Washington.

Figure 3 Short-Term Work and Layoffs for STC and Non-STC States



SOURCE: Calclulations based on Current Population Surveys, Outgoing Rotation Groups (all months) 1979–1993.

Table 6 Short-Time Work Rates for All Non-STC States and for Non-STC States with a Partial Unemployment Insurance Program by Level of Benefit Generosity[a] (%)

	As a share of employment				As a share of total hours			
	Level of benefit generosity for partial UI program				Level of benefit generosity for partial UI program			
Year	All non-STC states[b]	Least generous[c]	More generous[d]	Most generous[e]	All non-STC states[b]	Least generous[c]	More generous[d]	Most generous[e]
1979	1.09	1.18	1.09	0.85	0.35	0.38	0.35	0.30
1980	1.48	1.53	1.33	1.56	0.45	0.45	0.43	0.49
1981	1.44	1.50	1.35	1.37	0.44	0.44	0.43	0.46
1982	1.99	2.18	1.82	1.72	0.56	0.60	0.55	0.51
1983	1.45	1.54	1.37	1.31	0.43	0.46	0.41	0.45
1984	1.27	1.30	1.32	1.10	0.39	0.39	0.41	0.39
1985	1.20	1.26	1.23	0.99	0.37	0.36	0.39	0.37
1986	1.13	1.08	1.33	0.97	0.36	0.33	0.42	0.32
1987	1.11	1.10	1.11	1.15	0.35	0.35	0.34	0.36
1988	1.01	1.05	0.95	0.98	0.31	0.31	0.32	0.32
1989	1.01	1.08	0.92	0.96	0.32	0.30	0.33	0.35
1990	1.18	1.28	1.05	1.11	0.34	0.37	0.33	0.29
1991	1.39	1.50	1.30	1.22	0.44	0.48	0.40	0.41

(continued)

Table 6 (continued)

	As a share of employment				As a share of total hours			
	Level of benefit generosity for partial UI program				Level of benefit generosity for partial UI program			
Year	All non-STC states[b]	Least generous[c]	More generous[d]	Most generous[e]	All non-STC states[b]	Least generous[c]	More generous[d]	Most generous[e]
1992	1.10	1.11	1.21	0.93	0.33	0.33	0.35	0.31
1993	1.06	1.15	1.00	0.89	0.33	0.34	0.34	0.27

[a] The STW rates were calculated over all workers plus unemployed in each category using CPS Outgoing Rotation Group data for all months in the year.

[b] The "all non-STC states" category includes all states that have never had an STC program.

[c] The "least generous" category includes Alabama, D.C., Georgia, Indiana, Maine, Mississippi, New Hampshire, New Jersey, New Mexico, North Carolina, Ohio, Tennessee, and Virginia.

[d] The "more generous" category includes Colorado, Delaware, Hawaii, Idaho, Nebraska, North Dakota, Oklahoma, Pennsylvania, South Carolina, South Dakota, Utah, West Virginia, and Wyoming.

[e] The "most generous" category includes Alaska, Kentucky, Michigan, Montana, Nevada, and Wisconsin.

group taxes partial UI benefits at a rate of less than 100 percent for earnings over the disregard amount, providing the most generous potential benefits through partial UI. If the partial UI programs influence the use of STW, then STW should be most prevalent among those (non-STC) states that offer the most generous partial UI benefits.

An inspection of the numbers in Table 6 reveals that this is not the case: the employment incidence of STW is virtually the same for non-STC states that have the least generous partial UI programs compared to those that have the most generous partial UI programs. Even though the number of workers on STW appears unaffected by partial UI benefit generosity, there might still be greater hours adjustment due to incentives provided by partial UI schemes to dramatically reduce the length of the workweek. Yet total hours adjustment through STW is the same across the different levels of partial UI benefit generosity.

USING STW TO EVALUATE THE POTENTIAL IMPACT OF STC

The evidence presented here suggests that current patterns of STW in the United States appear to be primarily dictated by patterns of demand shocks and production technology, not by incentives provided through STC or partial UI. Thus the current patterns of STW and layoffs for each state should serve as a useful benchmark for researchers who wish to gauge the relative effect of proposed expansions in STC on STW and layoffs. However, doing this properly requires careful consideration of all possible factors that may affect STW, layoffs, or both, including parameters of state STC, regular UI, and partial UI programs. For example, it is possible that current STC programs have marginally increased the use of STW in the states that have introduced them. But a complete answer to this requires a determination of whether other incentives (and disincentives) to use UI and partial UI changed at the same time. Such an undertaking is beyond the scope of this study, but should be addressed in future research.

One question that can be answered is, where should we expect to see the highest rates of STC usage? The limited data on STC usage shows an uneven distribution across industries (Kerachsky et al. 1985;

Best and Mattesich 1980). This is not surprising given the patterns in Table 3, which show the highest rates of STW among the most cyclically and seasonally sensitive occupations. However, there is another potential explanation for the patterns in Table 3. Workers in whom the firm has invested the most training and/or who have the highest level of skills should be less likely to be put on STW, otherwise the worker on STW might take that as a negative signal about future employment at the firm and decide to look for a new job. To answer this it is necessary to disentangle industry shocks from occupational differences in the response to those shocks.

In order to sort out these effects, STW rates were calculated for six major occupation groups within the seven major industries (results not reported). Within each industry group, the more highly skilled management and professional occupations have the lowest rates of STW, which is consistent with firms wanting to protect investment in specific human capital. The clerical/administrative support occupation also has very low rates of STW, though always higher than managers. This probably reflects the fact that they embody less firm-specific human capital, yet work side-by-side with management, and so are slightly shielded from STW because of the direct support role they play to those workers least likely to be subject to STW.

The highest rates of STW within each industry are among the skilled and semiskilled laborers. The higher overall rate for laborers as a group, relative to white-collar workers, probably reflects both differences in production technology and levels of specific human capital. In particular, the higher rate of STW for semiskilled laborers compared to skilled laborers is probably due to their lower levels of skill.

STW rates tend to be comparable for the same occupation in different industries (for example, the STW rate for managers is similar across industries). However, the most cyclical industries exhibit the highest rates of STW for almost all occupations. For example, clerical/administrative support workers in construction have higher rates of STW than their counterparts in services, and the same holds for each of the other occupations within these two industries. The same is true for nondurable manufacturing compared to services. However, the pattern is less evident for durable manufacturing versus services.

These patterns suggest any expansion of existing STC programs should produce the greatest STC incidence in cyclical industries such as construction and nondurable manufacturing. Moreover, semiskilled laborers, skilled laborers, and, to a lesser extent, sales-related occupations should also have relatively high rates of STC usage, regardless of the industry of employment. Obviously, analyses such as these are only a crude first step at predicting STC take-up rates. A definitive answer to the effects of STC at the firm level requires firm-level data such as that used in the Mathematica evaluations (Kerachsky et al. 1985; Needels and Nicholson 1996). Yet STC predictions using CPS data and techniques such as those in this study should serve as a useful guide for researchers wishing to do more accurate analyses of the impact of STC than have been done to date.

CONCLUSION AND POLICY IMPLICATIONS

This study has documented short-time work patterns in the United States for 1968–1993. Despite very low STC take-up rates, STW is a prevalent phenomenon. The STW rate ranges from approximately 50 to 70 percent of the layoff rate.

The primary impact of STC program expansion most likely would be to subsidize those workers and firms that already use STW. The important implication of this is that vast numbers of workers could be put on the STC roles, thereby providing "evidence" that the programs were successful at averting layoffs, without impacting the incidence of layoffs at all. The key to determining the impact of STC program expansion on layoffs is not to count the number of people on STC alone, nor even to compare the number on STC relative to the number on layoff. Rather, layoff and STW rates—in terms of both workers and total hours—under STC must be compared to what they would have been in the absence of STC.

If the subsidy to STC is relatively large, additional workers will be put on STW relative to what would have happened otherwise, with a less than equal decrease in adjustment through layoffs. For example, if STC leads to 100 "additional" hours of STW, layoffs may be reduced by only 50 hours, with the additional 50 hours accounted for by an overadjustment through STW. The latter means a much greater distribution of the brunt of hours reduction across the workforce than is necessary.

In an era of tight budgets and reduced social welfare spending, this is a significant issue. The degree of imperfect experience rating of STC benefits is comparable to that for UI benefits. The extensive literature on excessive use of UI due to this subsidy suggests that widespread introduction and expansion of STC programs will lead to similar overuse of STC. Whether imperfect experience rating of both UI and STC benefits leads to a relatively greater overuse of layoffs, of STW, or of both is an empirical question. However, the overall net public subsidy to these two channels—layoffs and STW—would ensure an excessive impact of demand shocks on the existing pool of workers.

Unfortunately, the existing STC programs probably are too limited in scope to satisfactorily quantify the impact of the current STC system on employment versus hours adjustment in the United States. However, existing patterns of STW can be used to provide baseline estimates in future STC evaluations to determine that tradeoff. Similarly, differences in the relative subsidy to layoffs between states could be used to analyze how imperfect experience rating affects firms' choice of layoffs versus short-time work. Increased experience rating for regular unemployment insurance alone (as advocated by Hamermesh 1978, and Burdett and Wright 1989) may be sufficient to significantly tip the scales in favor of STW over layoffs, negating one of the primary arguments currently used by advocates for STC expansion.

Notes

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1. STC is frequently called "shared work" or "worksharing" by both researchers (Bednarzik 1980; Meltz and Reid 1983; Vasche 1982) and state UI agencies in the United States (National Foundation for Unemployment Compensation & Workers' Compensation 1996, p. 58). However, shared work is more commonly used to refer to permanent reductions in average hours per worker (Calmfors 1985;

Fitzroy 1981; Calmfors and Hoel 1989; Hart 1984; Riechel 1986), which is distinct from temporary reductions that are funded by partial UI benefits. The STC name is used exclusively in this chapter because it has no alternative interpretations.

2. Illinois discontinued its program in 1988 (Vroman 1990).
3. These figures have been adjusted for the bias in the pre-1994 Current Population Survey measures of STW and layoffs, detailed in the text.
4. In particular, no indication was given that the contracts were drawn from a representative sample.
5. Differences in overtime pay rules and related societal conventions may be particularly important unexplained factors not accounted for by Abraham and Houseman and by Van Audenrode. Their analyses treat increases and decreases in labor usage symmetrically, with no metric for measuring the difference between usual hours and actual hours worked. Thus much of cross-country differences and similarities that they measure may be identified by deviations above, not below, usual hours worked.
6. See note 5.
7. Bednarzik does not include workers on layoff when calculating the STW rate. Consequently, the rates reported here are not directly comparable to those in his study.
8. This was done by regressing usual hours on a host of demographic variables (race, marital status, age, education) and industry and occupation dummies separately for men and women for each year.
9. The seasonal pattern of STW in construction underscores the concern of STC program administrators that STC not be used to subsidize seasonal employment. However, despite this concern it is not clear whether existing STC guidelines are sufficient to prevent abuse by firms that experience predictable seasonal employment changes.
10. This estimate is perfectly accurate only if STW workers' mean usual hours are the same as non-STW workers' mean usual hours. This is probably a good approximation.
11. Illinois is not included in the Table 4 text because its program has been discontinued. However, Illinois is included as one of the STC states in all calculations because it did have an STC program during 1979–1993.
12. Note that the group of STC states includes observations for years in which some of the included states did not have an STC program (such as Kansas for 1979–1988 and Washington for 1979–1982).

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11

Economic Activities and the Demand for Work Sharing in Canada

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Work sharing refers to work-time arrangements in which all members of a work group reduce their hours of work to prevent the layoff of some members of the group. The reduction in working time and the associated reduction in income are thus redistributed over the entire work unit, rather than being concentrated on a few workers. In Canada this arrangement is formalized as the work sharing program under the authority of the Unemployment Insurance Act. Canadian work sharing was introduced as a pilot program in 1977 and was modeled on similar programs in effect in Europe. In 1981, in response to growing numbers of layoffs, the program was fully implemented and has been available since that time.

The work sharing program is based on the premise that it is better to keep workers employed than to have them experience a period of unemployment. Thus, its main objective is to maintain local, regional, and industrial employment levels during periods of short-term adverse economic conditions. Work sharing also has secondary objectives for both firms and employees. For firms, the program aims to assist them in retaining intact their skilled workforces and to help them avoid the costs associated with temporary layoffs such as recruiting and training new employees. For employees, the program aims to improve the level of income for workers who would otherwise be laid off and to assist workers in maintaining their skill levels and work motivation and reducing dislocation and uncertainty.

Under a work sharing agreement, layoffs are averted or postponed by reducing the workweek of employees in the designated core work group. An employer who intended to lay off 20 percent of employees for three months may use work sharing to reduce working hours of all employees by 20 percent over the same three-month period. Lost wages due to reductions of regular working hours are partially compensated by unemployment insurance (UI) benefits. Approximately 60 percent of lost wages are covered by UI benefits, charged to the UI Account. A UI-approved work sharing agreement is made between three parties: management, a majority of the affected workers, and Human Resources Development Canada (HRDC). Workers who apply for UI under work sharing do not have to serve the usual two-week waiting period for benefits. Employers must maintain fringe benefits for the duration of the work sharing agreement. Work sharing agreements may last for 26 weeks and may be extended to 38 weeks under special circumstances. A major evaluation of the program was completed in 1993, covering the years 1989 and 1990. The evaluation examined a sample of firms that used work sharing. Comparison firms were pre-screened and included in the sample if they had seriously considered laying off at least 20 percent of the members of one of their business units due to adverse economic circumstances. This procedure identified a sample of comparison firms who met the work sharing eligibility criteria but did not participate in work sharing.

The evaluation found that work sharing clearly avoids layoffs. However, in some cases layoffs may have been avoided without the work sharing program. Analysis of comparison firms showed that 7 percent of these firms did not lay off any employees. Further, in 29 percent of the work sharing cases, layoffs which should have been avoided by the program were merely postponed by the program, as these employees were laid off in the six months following program participation. Of these layoffs, 75 percent were of a permanent nature. Thus, in total, 64 percent of the layoffs that should have been averted by participation in work sharing can be said to have been avoided as a result of the program. Comparison with the 1984 evaluation of the same program found that the program was somewhat less successful in avoiding layoffs in 1989–1990 than it was in 1983.

The evaluation also found significant benefits of work sharing participation for workers. Participants who would have most likely suffered a layoff did much better than their comparison group counterparts who were laid off, experiencing a 19 percent reduction in income versus a 47 percent reduction of the layoff group. The work sharing group displayed much higher levels of morale, better attitudes to work and management, better social relations, and better physical and psychological health vis à vis those in the layoff situation.

Firms also experienced benefits from work sharing participation. They maintained the work sharing unit intact and expended $800–$1,800 less per layoff equivalent than comparison employers. They also returned to full production sooner than firms that laid off employees. However, there was no longer-term profitability or productivity advantage for these firms. This may suggest that work sharing is an appropriate tool to help firms deal with cyclical fluctuations in demand, but not for those firms facing fundamental, structural changes.

Work sharing was found to be more expensive for the UI fund than the layoff alternative. Costs were 33 percent higher for work sharing due to three factors: the waiver of the two-week UI waiting period for work sharers, the fact that 30 percent of layoffs never collected UI, and the incidence of layoffs in the post–work sharing period.[1] However, to balance these additional costs, there were also significant social benefits. The evaluation estimated that work sharing helped avoid costs related to the stress of unemployment, avoidance of costs related to unemployment scarring, and financial benefits to participating firms. Overall, the evaluation estimates a benefit-to-cost ratio of about 2.6:1.

While the 1993 evaluation answered many of the questions about the program, it did not examine the relationship between the potential demand for work sharing and economic conditions. For example, if the unemployment rates in the late 1980s were higher than the historical rates, what would have happened to the demand for UI-subsidized work sharing? What would have been the associated cost? Could the UI fund of the late 1980s have absorbed the additional cost?

This chapter concludes that the demand for work sharing is indeed sensitive to changes in economic activities. If the total unemployment rates were 1.37–1.50 percentage points higher in 1988–1990, then the demand for work sharing participation would have been 9 percentage points higher than it actually was in 1990. The estimated cost for the additional demand is $13.7 million. In relative terms, the worsening of economic conditions would have increased work sharing's share of total UI program expenditure in 1990 from 0.43 percent (actual) to 0.53 percent. Obviously, within a reasonable range of demand shocks, any increased demand for work sharing would have been too small to create a serious financing problem for the UI account.

THE LIKELIHOOD OF PARTICIPATION IN THE WORK SHARING PROGRAM

From our recent survey of the existing literature on work sharing, we have learned that the theoretical development in this field is primarily concerned with work sharing's effectiveness as a policy measure for alleviating the unemployment problem during recession.[2] In most of the theoretical discussions, the demand for work sharing has been taken as an inevitable phenomenon: during an economic downturn, some firms prefer to use work sharing to layoffs as an adjustment mechanism. Thus, other than relating demand for work sharing to business fluctuation, the existing literature is not very helpful in providing theoretical guidelines for specifying the demand for work sharing equation. In Canada, this problem is further complicated by the lack of useful time series for a comprehensive empirical investigation on this topic. Although the administrative files contain some aggregate time series on the number of work sharing applications approved, number of individuals in the program, and program expenditures, they have hardly any information on the behavior and characteristics of work sharing participating and nonparticipating firms. Thus, the idea of deriving a demand for work sharing equation from existing theories and estimating it directly from available data do not presently seem to be a feasible approach.

Recognizing the problems mentioned above, the quantitative work of this study circumvents the difficulty by working mainly with the cross-sectional data collected for Employment and Immigration Canada's 1993 evaluation.[3] Because these data have certain limitations, a specific methodology has to be developed to deal with them. The quantitative work includes four related components:

1) estimating a logistic equation that describes a firm's probability of participating in the work sharing program,

2) creating a relatively depressed scenario for the late 1980s (1987–1990) from a full-system econometric model simulation,

3) calculating the number of firms that would have become work sharing participants in the more depressed scenario, and

4) calculating the cost of the additional demand and its impact on the UI account.

The Logistic (Program Participation) Equation as a Demand Function

This study uses a logistic equation to estimate the probability of a typical firm that would participate in the work sharing program in 1990. The microdata are primarily from a special survey that Employment and Immigration Canada (EIC) used to conduct its 1993 evaluation on the work sharing program. The data consist of 310 participating and 256 nonparticipating firms in 1990.[4] The participating employer sample was selected from an administrative file that contained the names of work sharing firms during 1990. Members of the comparison group sample were selected from the EIC Record of Employment file; these firms were selected on the basis of their comparability to work sharing firms in characteristics and activity experience[5] (e.g., members of the comparison group must have laid off workers in 1990). In carrying out the econometric estimation, the procedure requires full information (no missing data points) for the dependent and independent variables. Because of missing values for selected variables, 43 firms have to be excluded from the sample. Thus, the final econometric estimation is based on the information from 289 participating firms and 234 members of the group of comparable employers.

Ideally output or sales data for periods immediately prior to the firm's applying for work sharing participation (program participants) or laying off workers (members of the comparison group) should be used as a measure of the firm's business fluctuation. Unfortunately, the survey was not designed to deal with the demand issue and did not collect any output or sales data. This forces us to search for a suitable proxy. The proxy used is the unemployment rate of the UI region where the firm is located. This variable is chosen for two reasons. First, even though the UI regional unemployment rate, by definition, cannot claim to be a unique economic activity indicator of the firm, it is specific to the economic climate where the firm operates. Second, the time series for this variable is available. The data are from Statistics Canada's Labour Force Survey unpublished worksheets. We have the UI region data on labor force, employment, unemployment, and unemployment rate dating back to 1979. The availability of these time series is operationally very important because it allows us to introduce a dynamic element into the specification of the equation specification (see the discussion below).

Even if one accepts the UI regional unemployment rate as a reasonable proxy for approximating economic downturns, proper timing and functional form remain crucial to the specification of the probability of program participation equation. First, using the regional unemployment rate of 1990 as an explanatory variable would present a serious technical problem. In 1990, the participating firms were already in the work sharing program. If the program was effective in lowering unemployment, then the UI regional unemployment rate of 1990 would also be dependent upon the extent of program participation within the UI region. This simultaneity bias presents an interpretation problem because the estimated equation would have mixed the program's effects on regional unemployment with the influence of economic activities on the demand for work sharing. Second, if a firm uses UI-subsidized work sharing as an adjustment mechanism for the decline in the demand for its products, then the proxy variable should probably be in the first difference form rather than level form. This line of reasoning suggests that the proper variable for explaining a firm's probability for program participation should be the change in the unemployment experience of the UI region prior to the firm's decision to join (or not to join) the work sharing program.[6] This specification would not have been operational if the time series for the UI regional unemployment rate is not available. Symbolically, the specification for the program participation equation may be summarized as follows: the probability of firm *i* participating in the work sharing program is

$$\text{Probability}_i = f(\Delta\text{URATE}_{i,t-1}, X_{i,t}).$$

$\Delta\text{URATE}_{i,t-1} = \text{URATE}_{i,t-1} - \text{URATE}_{i,t-2}$, URATE_i denotes the unemployment rate of the UI region where firm i is located, t refers to the current period, and $t - 1$ the period prior to participating in the work sharing program. X_i represents a vector of firm specific attributes, including the average skill rating of the firm's employees, percentage of employees unionized, organization structure (whether the firm operated at one single location, multiple locations in one province, multiple locations across Canada, or multiple locations internationally), the firm's industrial affiliation, the firm's years of operation, etc.

The logistic equation, when estimated, serves certain purposes. First, it tests the hypothesis that the demand for work sharing depends on changes in economic activities. If the estimated coefficient for the $\Delta\text{URATE}_{i,t-1}$ variable is positive and statistically significant, it would confirm that more firms would like to become program participants as the economic climate worsens. Second, it provides the reader with some information on which other exogenous forces influence a firm's program participation decision. Third, the estimated equation provides us a means of calculating a firm's probability of participating in the work sharing program under various unemployment conditions. This last function is crucial to the theme of this study and will later become apparent. The estimated coefficients and essential statistics are shown in Table 1.

In addition to the explanatory variables listed above, earlier versions of the estimated equation also included type of organizations (private sector, public or nonprofit organizations), number of full-time workers employed by the firm, and provincial dummies on the right-hand side of the equation. They were subsequently dropped for various reasons. The effects of provincial differences were conceptually and empirically reflected in the recent changes in the UI region unemployment rate variable.[7] Therefore, it was not necessary to include provincial dummies as additional explanatory variables. The other variables were excluded because they were statistically insignificant and their exclusion did not noticeably affect the estimated coefficients of the other explanatory variables.

Although our main interest is in the "recent changes in the UI region unemployment rate" variable, the estimated coefficients of other

Table 1 Estimated Logistic Equation

Dependent variable: employer program participation (1=Yes, 0=No)
Number of firms included in the analysis: 523

–2Log likelihood	651.791		
Goodness of fit	525.196		
	X^2	df	Significance
Model X^2	67.446	12	0.0000
Improvement	67.446	12	0.0000

Variable	Coefficient	Standard error	Wald	Significance
Employee's average skill rating	0.2122	0.0710	8.9247	0.0028
Recent change in the UI region unemployment rate	0.2777	0.1370	4.1089	0.0427
Organization operations				
Single location	0.2517	0.2907	0.7499	0.3865
Multiple (across Canada)	0.0630	0.4005	0.0247	0.8750
Multiple (international)	–1.1829	0.5089	5.4041	0.0201
Multiple (one province)	0.0000			
Type of industry				
Primary	–0.9615	0.5581	2.9682	0.0849
Heavy manufacturing	0.1080	0.2558	0.1784	0.6728
Construction	–1.3049	0.3497	13.9260	0.0002
Trade	0.4674	0.2830	2.7285	0.0986
Other	0.4110	0.3381	1.4778	0.2241
Light manufacturing	0.0000			
Percentage of employees unionized	–0.0036	0.0019	3.5111	0.0610
Years of operation	0.0079	0.0046	2.9362	0.0866
Constant	–0.9074	0.4469	4.1226	0.0423

explanatory variables are also of some relevance. Since the estimated coefficients of the logistic equation cannot directly tell us the effects of the explanatory variables on a firm's probability of program participation, as an illustrative example we have calculated the possible impact of each explanatory variable on participation probability by holding all other explanatory variables constant at specific values. Appendix A reports the results.

"Employee's average skill rating" significantly influences a firm's probability of using the work sharing program.[8] All other things being equal, a firm with many highly skilled workers tends to use the UI-subsidized work sharing as the demand adjustment mechanism more often than firms that employ a relatively large number of unskilled workers. This is consistent with the notion that the option of work sharing participation rests mostly with employers. It is the cost-minimization conditions that determine this behavior: Laid-off workers of relatively high skills are more likely not available for subsequent rehiring when the firm's business starts to pick up; training new workers to fill these positions would be a relatively costly option to the firm.

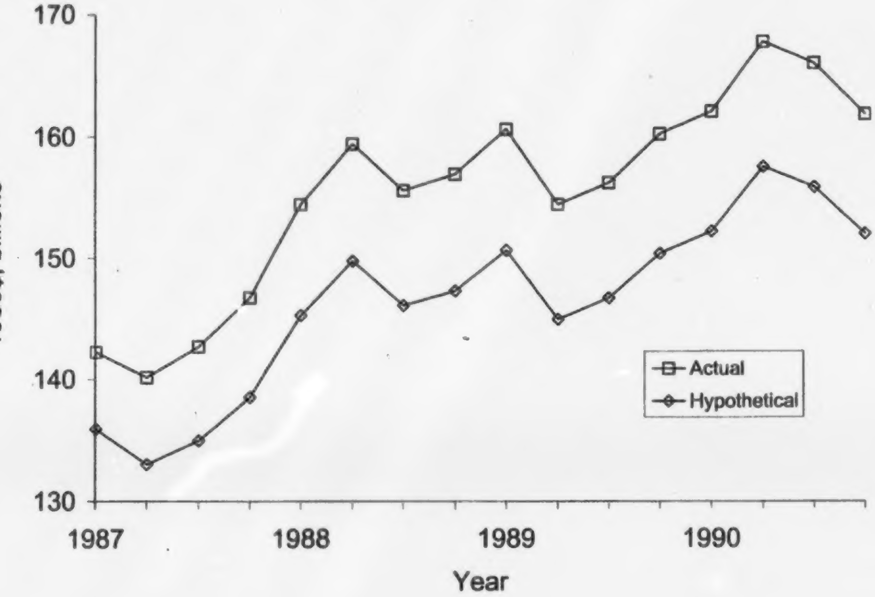
The age of a firm's establishment (in terms of their years of operation) also seems to have a positive influence on the firm's program participation decision, but this result is not statistically persuasive. (The estimated coefficient is statistically significant at the 10 percent level but not at the 5 percent level.) How long the firm has been in business should not greatly affect its present profit-maximization (or cost-minimization) conditions. Thus, the estimated coefficient of this variable is meaningful only if it is an approximation of the firm's outlook of future business prospect. In other words, a more established firm tends to be more optimistic of its future than the relatively new companies. Therefore, it is more willing to use work sharing to maintain its labor force during business slow-downs. The unionization of workers exerts a negative influence on the firm's participation probability. The statistical result, however, is not as strong as expected. While most of the local unions prefer work sharing to worker layoffs during economic downturns, very few centralized unions endorse the work sharing option because it erodes the seniority principle. These two opposite forces are probably sufficient to prevent this variable from becoming statistically very strong. The organization and industry dummies present a mixed bag of results. Some are highly significant and some are not significant at all. As a principle, we keep all of them in the estimated equation, even though dropping the insignificant dummies would not have noticeably affected the rest of the estimated coefficients.

The estimated coefficient for the "recent change in the UI region unemployment rate" variable is 0.2777 and is statistically significant at the 4 percent level. Although the estimated coefficient corresponds to the dependent variable in a log (odds) form[9] and therefore cannot directly tell us the impact of this variable on the probability of program participation, the positive coefficient confirms our *a priori* expectation. Later in this section we will provide some impact estimates based on the estimated equation and simulation techniques.

Different Economic Scenarios, 1987–1990

The logistic equation by itself is still not capable of answering the questions posed in the introductory section of this paper. For example, the estimated logistic equation would show that if the expansion phase of the business cycle of the 1980s ended earlier than it did, more firms would probably have wanted to join the work sharing program. This does not give us a quantitative estimate of the size of the additional demand that could have resulted from a more depressed economic climate in the late 1980s. Apparently, one cannot obtain such a quantitative estimate without specifying the deterioration of economic activities in quantitative terms. The simplest way to meet this information requirement is to assume that the total unemployment rates in the late 1980s were higher than their historical counterparts by certain percentage points. These figures can then be distributed proportionally to the UI regions to yield a set of hypothetical UI regional unemployment rates, which can in turn be fed into the logistic equation for further investigation. In this study, we prefer a more plausible hypothetical scenario than the arbitrarily assumed one. The hypothetical (more depressed) scenario used in this study is from the solution of a full-system econometric model,[10] in which Canadian exports, including automobiles and parts but excluding other manufactured goods and mining products,[11] in 1987–1990 were assumed to be 10 percent less than they actually were historically (see Figure 1). In this hypothetical setting, because of the decline in aggregate demand resulting from the assumed

Figure 1 Canadian Exports, 1987–1990

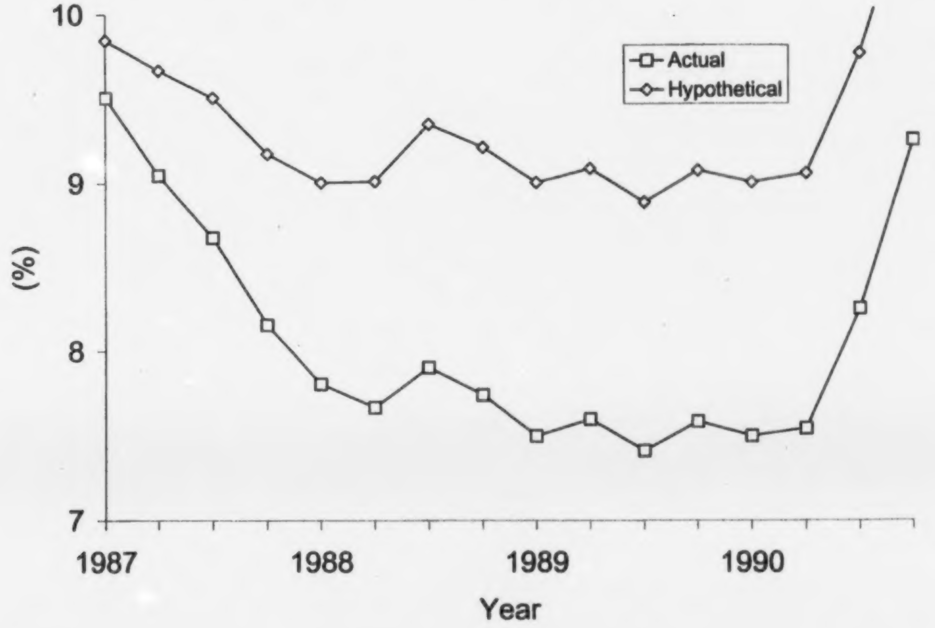


drop in exports, more individuals were expected to become unemployed. The additional unemployed individuals can then be distributed to the UI regions according to their labor force shares. From these new unemployment figures we may calculate the UI regional unemployment rates for the hypothetical scenario.

The end result is that, in the hypothetical scenario, more individuals would have been unemployed and the impacts on the total unemployment rates were noticeable (0.70, 1.37, 1.49, and 1.50 percentage points higher than the historical figures in 1987, 1988, 1989, and 1990, respectively). The unemployment rates for the actual and hypothetical scenarios are graphically presented in Figure 2.[12] Furthermore, we have distributed the additional unemployed individuals (not shown here but available in the solutions of the model simulations) of the hypothetical scenario across 49 UI regions, according to the labor shares of the UI regions, and recalculated the UI regional unemployment rates for the more depressed (hypothetical) scenario.

Although this macrosimulation is not essential to the quantitative work of this study, we have decided to use it. The model solution generates a reasonably realistic but more depressed economy than the

Figure 2 The Unemployment Rate, 1987–1990



actual experience of 1987–1990. It also illustrates that exogenous forces could have easily ended the expansion phase of the 1980s business cycle much earlier than it did.

Expected Number of Participating Firms under Different Economic Scenarios

The estimated logistic equation, the actual UI regional unemployment rates, and the UI regional unemployment rates for the hypothetical scenario provide us with the required tools and information for calculating the probability of program participation for each firm for two scenarios (base-case and the hypothetical). First, for the base-case, we obtain a set of probability estimates for all firms (including participants and members of the comparison group) by inserting the actual values of all explanatory variables into the unscrambled logistic equation.[13] Similarly, for the hypothetical scenario, by replacing the actual UI regional unemployment rates with their hypothetical counterparts while keeping the actual values of other explanatory variables unchanged, we may calculate a set of estimates for the firms' chances of participating in the work sharing program under the more depressed economic climate. The first set of estimated figures shows each firm's probability of participation, with the values of all explanatory variables identical to their actual (historical) values. This may be labelled as the base-case probability. The second set is similar to the first, except that the calculation replaces the actual UI regional unemployment rates with the hypothetical scenario's UI regional unemployment rates. In other words, the second set shows each firm's probability of participation under the more depressed economic conditions of the hypothetical scenario, while holding all other things constant.

The estimated probability provides us with the information concerning a firm's chance of becoming a program participant, but it still does not tell us whether or not the firm would indeed be in the program. After all, even a firm with a probability of 90 percent participation still has a slim chance of not being a participant. In this study we use the random-draw simulation technique to determine whether a firm is *in* or *out* of the work sharing program. The procedure is identical to drawing a "chip" randomly from a hat. For example, the participation probability for a certain firm was usually 70 percent (the base-case), but under the more depressed economic climate of the hypothetical scenario its probability increased to 71 percent in 1990. To determine whether or not this firm would become a program participant in the base-case and in the hypothetical scenario, we create two separate hats. The first hat would have 70 chips marked "in" and 30 marked "out," while the second hat would have 71 chips marked "in" and 29 chips marked "out" to reflect its slightly higher probability of program participation. We would then randomly draw one chip from each of the hats and record the results of the random draws. Repeat the same procedures for all firms in the sample. The difference between the total numbers of "in" firms in the two scenarios (base-case and hypothetical) would be taken as the estimated impacts of the more depressed economic climate on the demand for work sharing. This is, however, only the result of one random-draw experiment. The results in Table 2 are the averages of 10 experiments.[14]

Table 2 Results of Random-Draw Simulations (based on a sample of 523 firms)

Total number of participating firms (base-case)	Total number of participating firms (hypothetical)	Difference (hypothetical/base-case)
277	302	25

The simulated figure for the base-case underestimates the actual number of participants in 1990 by 12 firms (an error of 4 percent). Since the estimated logistic equation cannot be expected to predict the probability of participation perfectly and the random-draw experiments have been conducted only 10 times, the goodness of fit appears to be acceptably close.

The simulation results suggest that under the influence of a worse economic climate, as specified by the hypothetical scenario, the demand for work sharing participation would have been 9 percentage points higher than it was in 1990. Since the results reported are based on a sample of 289 participating and 234 nonparticipating firms, we have to mark up the total number of participating firms (work sharing applications approved) in 1990 by 9 percent to yield an estimate of participating firms for the total economy; that is,

hypothetical scenario: number of participating firms, total economy, 1990 = 6,297 × (1 + 0.09) = 6,873,

where 6,297 is the actual total number of work sharing applications approved in 1990. In other words, the more depressed economic climate of the hypothetical scenario would have induced 576 more firms to participate in the work sharing program in 1990. This estimate should, however, be taken as an illustrative example rather than a definitive answer. First, aside from the imperfection of the econometric and simulation techniques, the analysis is based on a relatively small sample size and the data were not originally collected to test the sensitivity of work sharing demand to economic activity fluctuation. In the future, evaluators should probably take the demand dimension as an integral part of the evaluation framework and revisit this topic. Second, the estimate depends directly on the degree of activity slowdowns created by the macromodel simulation. The deterioration outlined in the hypothetical economy is only one of the many plausible scenarios. A different hypothetical scenario would, of course, yield different results.

ASSOCIATED COSTS, IMPACTS ON THE UI ACCOUNT, AND IMPLICATIONS

There are at least three remaining questions that we should attempt to answer: 1) What is the cost of the additional demand for work sharing? 2) What is its impact on the UI Account? 3) How would the government have reacted to the additional applications for the work sharing program?

Costs and Impacts on the UI Account

From the administrative data, we know that a participating firm in 1990 cost the government an average of \$9,798.[15] However, this would be, *a priori*, a downward biased estimate for calculating the associated costs for our purposes. It fails to account for the additional utilization of work sharing among firms already in the work sharing program when unemployment increases. The available data do not allow us to calculate this downward bias accurately. In this study, we use the 1990 administrative data to approximate the relationship between UI regional work sharing expenditure and the UI regional unemployment rate (weighted by the region's employment share). The estimated equation suggests that the average cost for a participating firm would have been about 11.9 percent higher than the actual average cost in the hypothetical scenario.[16] Based on this information, we may approximate the cost of the additional demand for program participation in the hypothetical scenario as follows:

(i) Estimated cost of additional demand = increased cost for firms already in the work sharing program in 1990 + cost for financing 576 additional participating firms resulting from the worsening of economic conditions in the hypothetical scenario
(\$9,798 × 0.119 × 6,297) + (\$9,798 × 1.119 × 576)
= \$13.66 million.

(ii) Actual work sharing program expenditure in 1990 = $61.7 million.
(iii) Total UI expenditure in 1990 = $14,355 million.
Ratio A = 100 × (i)/(iii) = 0.095%.
Ratio B = 100 × (ii)/(iii) = 0.429%
Ratio C = 100 × [(i)+(ii)]/(iii) = 0.525%.

In 1990, the total Unemployment Insurance Developmental Uses (UIDU) were substantially below the maximum of 15 percent allowed by law (Bill C-21). Work sharing expenditure in this year accounted for less than 18 percent of the total UIDU expenditure. These statistics, along with the fact that work sharing was a relatively small program option,[17] suggest that the government could have easily absorbed the additional demand for work sharing of the hypothetical scenario by a minor reallocation of UI funds while keeping the total UI program expenditure of 1990 unchanged.

Government's Response to the Demand for Work Sharing

Figure 3 shows the relationship between the unemployment rate and work sharing applications approved. The correlation between them is positive but statistically insignificant. (The simple correlation coefficient for the variables equals 0.53, which is not even statistically significant at the 10 percent level.[18]) Figure 4 presents the graph for the change in the unemployment rate[19] and work sharing applications approved. It becomes obvious that the two variables are closely correlated with each other. The simple correlation coefficient is 0.90, which is statistically highly significant. As contended earlier, firms' demand for work sharing is related to the change in economic activities rather than the level of activities. In this examination of the aggregate time series, we have found other indirect, circumstantial evidence to support this contention.

If the "work sharing applications approved" series is interpreted as the locus of the equilibrium points between the demand for and supply of work sharing with the government adopting a 100 percent accommodative policy,[20] then the data should reflect the demand and supply information equally well. The existing data seem to suggest that government's policy has been quite "accommodative." In 1982–1983, the

Figure 3 The Unemployment Rate and Work Sharing Applications Approved

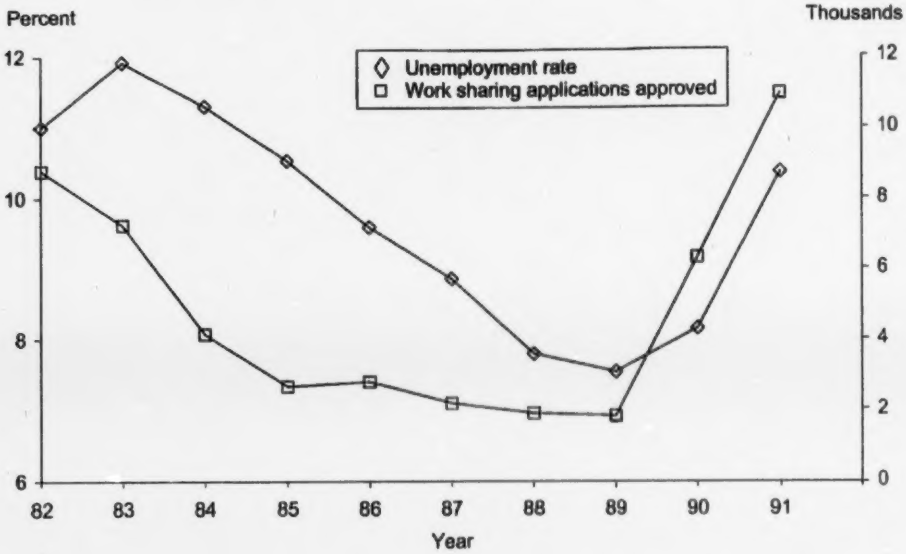
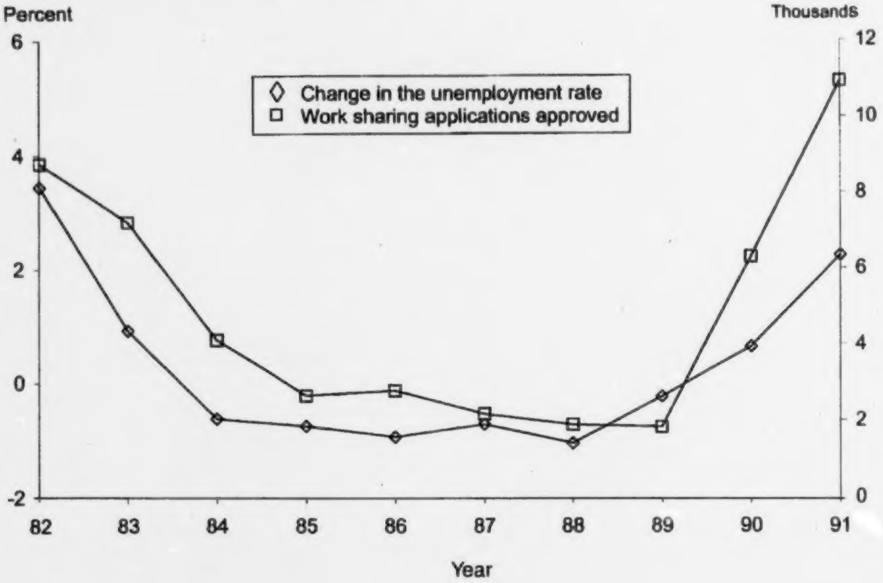


Figure 4 Change in the Unemployment Rate and Work Sharing Applications Approved



average change in the unemployment rate was 2.18 percent; the average number of applications approved was 8,009 per annum. From 1984 through 1989, a period of uninterrupted economic expansion, the average change in the unemployment rate was –0.733 percent, and the annual average of applications approved declined to 2,573 firms. In 1990–1991, the change in the unemployment rate became positive again (the average change was 1.42 percent), and the number of work sharing applications climbed to 8,613 per annum. This sensitivity to changes in the unemployment rate suggests that had the economic climate in the late 1980s become worse than it actually was, the government could have probably absorbed the additional demand.

Reid and Meltz (1983) and Pal (1983) note that the Canadian government's interest in work sharing has risen and fallen with changes in the unemployment rate. They argue that instead of implementing policy on the basis of careful long-run planning, the use of the program as an ad hoc response to the crisis of rising unemployment obviously leaves a lot to be desired. Their observation on the sensitivity of work sharing applications approved to changes in the unemployment rate has been quite accurate. In recent years, the sensitivity seems even higher. However, their criticism of the government's accommodative approach may have been too harsh. One would expect that the work sharing program, especially work sharing as a passive policy measure,[21] should always be responsive to the demand of firms. Whether or not the program could have played a more active role in the Canadian labor market is a moot question. Not only has France's experience of using work sharing as an active policy (job creation) not been convincingly successful, the relatively small size of the Canadian work sharing program does not suggest that it has the potential of creating a large number of jobs. Given the existing fiscal stance of the government, expanding the program for the sake of testing out the effectiveness of work sharing as a job creation policy must be rated as one of the most unlikely events in the foreseeable future.

CONCLUDING OBSERVATIONS

In the last decade in North America, corporate restructuring has imposed some alternatives to conventional working time arrangements. The increased use of nonstandard forms of work, including part-time, contract, and outsourcing, is associated with such restructuring in the context of competitive cost reduction. At the same time, the Canadian unemployment rate increased dramatically to a 9.5 percent average in the 1980s and the 1990s.

Policy responses to the growth of nonstandard work are now just emerging, starting with the changes in the Canadian employment insurance scheme to make eligibility for benefits based on hours, not weeks. An hours-based system better reflects current work patterns, particularly the rise in part-time and multiple job hours.

The other major insurance policy response is the work sharing program, which, in its current design, is clearly a countercyclical measure to enable firms to hoard labor or for workers to share unemployment during downturns. In its design, the program is not available to subsidize corporate restructuring. Two formal evaluations of this program have shown that work sharing does make a difference in averting layoffs, and, despite being more expensive to the UI account than straight layoff benefits, the economic and social benefits accruing to participating firms and workers more than offset the program costs. This chapter extends the evaluation work by looking specifically at the relationship between changes in economic activities and the demand for work sharing.

To examine this relationship, this study uses microsimulation as well as macrosimulation techniques. In a full-system macrosimulation, a 10 percent reduction of Canadian exports of nonmanufacturing/nonmining products and exports of automobiles and auto parts in 1987–1990 would result in a 1.5 percentage point increase in the unemployment rate by 1990. This change in declining economic activity would increase work sharing participation by 9 percent, at a cost of an additional $13.66 million in 1990 to a $14.35 billion UI account for that year. Within the existing legislative and regulatory framework, such an increased demand of work sharing could easily have been absorbed.

The 1993 evaluation results suggest that the additional expenditures could have been cost-effective as well.

This finding, of course, is relevant to the program as it currently operates. It sheds little light on how work sharing might be extended under current rules or how it might be used in an aggressive redesign of working time or under different rules for active job creation purposes. Suggestions have been made by both policymakers and academics that work sharing agreements might reduce UI premiums to firms that create and finance new jobs to compensate for the reduction in working time of designated employees. Under these circumstances, work sharing may create job opportunities for youth and other unemployed groups back-filling designated positions. This chapter does not address this policy debate directly. What it shows clearly is that work sharing is sensitive to the change rather than the level of economic activity, and that the probability of work sharing participation is higher among firms with higher-level skilled workforces.

Job creation stimulation is more common at the entry skills level. This suggests that the present program is limited in its potential as a job creation initiative. Even if work sharing is a good investment, as the evaluation results show, the cost of providing it to all potential layoff situations may be prohibitive. This chapter shows that an increased demand for work sharing can be accommodated as a relatively small program option under UIDU. In the current fiscal environment, it is difficult to imagine a proactive use of work sharing as a job creation mechanism without finding new monies or at least reprofiling UIDU expenditures at the expense of the other two major UIDU activities: UI-sponsored training and job creation partnerships. Finally, this chapter raises some questions about the appropriateness of work sharing as a job creation stimulus directed at firms that are primarily interested in maintaining a skilled workforce. Policymakers would need to take this present feature of work sharing participation into consideration if they were to redesign work sharing as both a job maintenance and job creation program. Before embarking upon this, however, it would be instructive to study more closely the work sharing experience in France in the 1980s and the reasons why the program reverted back from a job creation initiative to an employment maintenance scheme.

Notes

We are grateful to Garnett Picot of Statistics Canada and Wayne Vroman of Urban Institute for their comments on various aspects of the work contained here. All errors and omissions remain, of course, our responsibility.

1. The 1993 evaluation study might have overlooked two other factors that could have contributed to the comparatively high cost of the work sharing program. First, relative to the layoff option, work sharing tended to include a higher share of senior workers. Since earnings and seniority are positively correlated, the average work sharing benefit would be higher than the average regular UI benefit. Second, work sharing would be more expensive than regular UI payments if the layoff conversion ratio was less than unity, i.e., the increase in work sharing weeks exceeded the decrease in layoff weeks.
2. See, for example, MaCoy and Morand (1984), and Owen (1989).
3. See Employment and Immigration Canada (1993).
4. The survey also contains information for the year 1989. Because of the time constraint, we have decided not to duplicate the empirical work for 1989. In the early stages of an economic downturn, firms are not sure whether dismissals and long-term layoffs are necessary; their demand for work sharing may be different from those at different points of the business cycle. Future work on this topic should probably investigate the sensitivity of work sharing demand at different points of the business cycle as well.
5. For a detailed description of survey design and characteristics of participating firms and members of the comparison group, see Employment and Immigration Canada (1993).
6. Empirically, when the level of regional unemployment rate enters the right-hand side of the equation, the coefficient is positive but statistically insignificant. On the other hand, the first difference of the regional unemployment rate is statistically significant, and the result seems quite robust. The inclusion or exclusion of other explanatory variables in the specification does not significantly alter the result.
7. This variable is defined as ($URATE_{i,1989} - URATE_{i,1988}$); $URATE_i$ denotes the unemployment rate of the UI region where firm *i* is located, and the second subscript refers to a specific year. When the variable is expressed in terms of relative change, [i.e., ($URATE_{i,1989} - URATE_{i,1988})/URATE_{i,1988}$], the estimated coefficient remains positive but the statistical result is substantially weakened. It is statistically different from zero only at the 12.6 percent level. The statistical results for the other estimated coefficients remain basically the same. This finding suggests that workers and employers take the absolute change in the unemployment rate as a change in economic conditions, but they are probably unfamiliar with the concept of the relative change in the unemployment rate.
8. In the survey, the employer was asked to rate the firm's employees' literary skills, numeracy skills, and technological literacy separately, with a rating of 1 denoting

extremely low in the category, 4 average, and 7 extremely high. The variable used in the logistic regression equation is the average of the three skill variables.

9. That is, ln $[p/(1-p)]$, where p denotes the probability of work sharing participation.
10. The econometric model consists of about 300 behavioral equations and identities. It is a modified and extended version of the Conference Board's PC–Canadian Model (PCCDN). See Conference Board of Canada (1989).
11. The exports of other manufacturing goods and mining products are endogenously determined in the full-system econometric model.
12. The difference between the base-case and the hypothetical scenario represents the impact of the assumption on total Canadian exports. In this report, for the sake of simplification and interpretation convenience, we add this difference to the actual data of the variable in question. This procedure allows us to compare the hypothetical scenario figures directly to the historical data in level form.
13. The dependent variable for the estimated logistic equation is in the form of $\ln[p/(1-p)]$, where p denotes the probability of program participation. Feeding the values of the explanatory variables directly into the estimated equation would give us the ln odds rather than the probability.
14. The random draw process is such that a participating firm may or may not be classified as "in" the work sharing program; similarly, a nonparticipant is not necessarily out of the program in the experiment. Since the final tabulation compares the simulated figures of the two scenarios (base-case and hypothetical), this lack of perfect fit should not present an interpretation problem. The errors are random and should be cancelled out in the process of calculating the differences.
15. This figure, which is the ratio of the total work sharing expenditures in 1990, to the number of applications approved in 1990, is only the short-run book value. It does not take into account the addition of UI benefits paid to work sharing workers who were laid off in the post–work sharing period. See Employment and Immigration Canada (1993).
16. The estimated equation:

$$\text{Work Sharing Expenditure}_i = \underset{(0.76)}{180.95} + \underset{(5.31)}{4170.22} \times \text{URate}_i \times \text{EShare}_i,$$

$$\text{adjusted } R^2 = 0.37,$$

where a) figures in parentheses are t-statistics; b) i: UI region i; URate_i: the unemployment rate of UI region i; Eshare_i: the employment share of UI region i (i.e., employment of UI region i/total employment).

17. Why did so few employers use the work sharing program as a means for adjusting workers hours? The answer to this question is not obvious, and is a research topic itself. Work sharing has always been a relatively small program in Canada. From 1982–1995, the total unemployment rate fluctuated between 6.2 and 10.7 percent, the ratio of "work sharing weeks paid to total UI benefits weeks paid" remained in the "0.3 to 2.9 percent" neighborhood.
18. We have only 10 observations for this calculation.

19. That is, the first difference of the unemployment rate, $URATE_t - URATE_{t-1}$.
20. This assumes that the government usually approves all legitimate applications for work sharing participation.
21. For a discussion on active and passive work sharing, see Tremblay (1989).

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Table A1 Estimated Effects of Explanatory Variables on Program Participation Probability[a]

Variable	Probability
Employee's average skill rating	0.052
Recent change in the UI region unemployment rate	0.068
Organization operations	
Single location	0.000
Multiple (across Canada)	0.000
Multiple (international)	-0.272
Multiple (one province) – reference case	0.000
Type of industry	
Primary	–0.228
Heavy manufacturing	0.000
Construction	–0.295
Trade	0.113
Other	0.000
Light manufacturing – reference case	0.000
Percentage of employees unionized	–0.001
Years of operation	0.002

[a] Each value denotes the marginal effect of one additional unit of the variable on a typical firm's program participation probability, evaluating at the mean of the variable and holding all other explanatory variables constant at their mean values. If the variable is a 0 or 1 dummy variable, then 0 is taken as the mean in the calculation. "Organization" and "industry" variables statistically insignificant at the 10 percent level are treated the same as their respective reference cases.

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